Effects of the Changing U.S. Age Distribution on Macroeconomic Equations

By RAY C. FAIR AND KATHRYN M. DOMINGUEZ*

The effects of the changing U.S. age distribution on various macroeconomic equations are examined in this paper. The equations include consumption, housing-investment, money-demand, and labor-force-participation equations. There seems to be enough variance in the age-distribution data to allow reasonably precise estimates of the effects of the age distribution on the macro variables. (*JEL* 131)

A striking feature of postwar U.S. society has been the baby boom of the late 1940's and the 1950's and the subsequent falling off of the birth rate in the 1960's. Figure 1 shows a plot of the number of births by year for the period 1900–1986. The number of births rose dramatically from 2.5 million in 1945 to 4.2 million in 1961 and then fell back to 3.1 million in 1974. Figure 2 shows the consequences of this birth pattern for the percentage of prime-age (25–54) people in the working-age (≥ 16) population. In 1952 this percentage was 57.9, whereas by 1977 it had fallen to 49.5. Since 1980 the percentage of prime-aged workers has risen sharply as the baby boomers have begun to pass the age of 25.

In this paper, we examine the effects of the dramatic changes in the U.S. population age distribution on the behavior of several macroeconomic variables. Previous empirical studies of aggregate behavior have typically ignored age-distribution effects, relying instead on the representative-agent paradigm. While this approach is attractive in that it derives macroeconomic relationships from a well-specified individual's optimizing problem, it explicitly assumes that the population is homogeneous. Thomas M. Stoker (1986) shows that, in order for representative-agent models to describe accurately aggregate behavior, all marginal reactions of individuals to changes in aggregate variables must be identical. This condition seems particularly unlikely to hold across individuals from different age cohorts.

Here, we abandon the representative-agent approach in favor of models that explicitly allow for age-based population heterogeneity.[1] In Section I, we test the

*Fair: Cowles Foundation for Research in Economics, Yale University, New Haven, CT 06520, and National Bureau of Economic Research; Dominguez: Kennedy School of Government, Harvard University, Cambridge, MA 02138, and National Bureau of Economic Research. We are indebted to the referees for many helpful comments.

[1]Previous studies using a similar modeling approach have used both population and income data to examine the influence of distributional changes on interest rates, saving behavior, and commodity expenditures. Henry McMillan and Jerome Baesel (1987) find that real interest rates are negatively related to the ratio of savers to borrowers in the postwar United States. Their simulation results predict negative real interest rates beginning in the 1990's as baby boomers reach the saving phase of their life cycle. Charles Lieberman and Paul Wachtel (1980) find that the saving rate in the 1970's would have been substantially smaller if the demographic structure that existed in the 1960's had remained unchanged. Dale Heien (1972) includes the median age of the population over 24 in his life-cycle study and finds that the saving rate has increased with the median age of the adult population. Frank Denton and Byron Spencer (1976), however, using both Canadian and U.N. census data, find no age-distribution influence on aggregate saving and consumption.

Alan S. Blinder (1975) tests for the effect of the income distribution on saving behavior and reports no systematic relationship. Blinder concludes, however, that this result may stem, not from the lack of a relationship, but rather, from the relative stability of the U.S. income distribution since World War II. Stoker (1986) finds a significant relationship between income distribution and commodity expenditure, although he

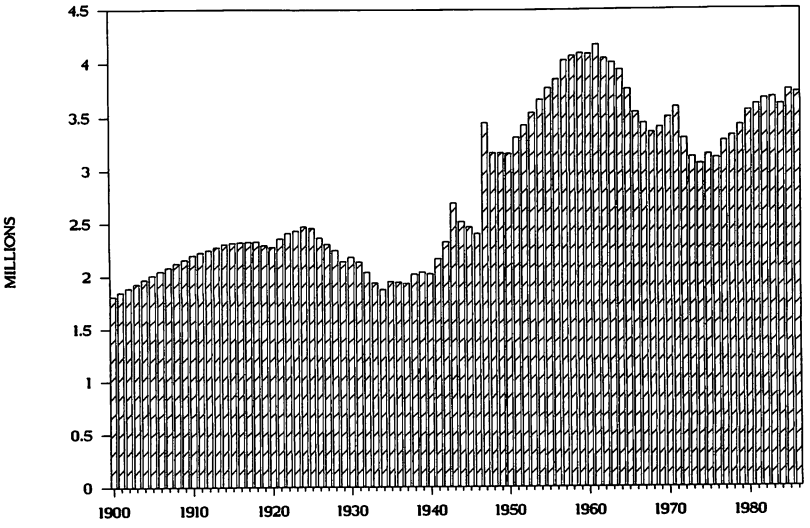

FIGURE 1. NUMBER OF BIRTHS BY YEAR, 1900–1986

Source: Current Population Reports, Series P-25, numbers 310, 519, 917, 965, and 985 (U.S. Department of Commerce, Bureau of the Census).

hypothesis that the changing age distribution is significant in explaining aggregate consumption and housing investment. The life-cycle model of Franco Modigliani and Richard Brumberg (1954) and the extension by Albert Ando and Modigliani (1963) predict that people in their prime working years consume a smaller fraction of their current income than do people younger and older. Further, if housing consumption is roughly proportional to the stock of housing, then the life-cycle model implies that the housing stock relative to current income should likewise vary with age. Prime-age people should consume less housing relative to their income than those younger or older.

The money-demand models of W. J. Baumol (1952) and James Tobin (1956) predict that there is a positive relationship between the transactions costs associated with obtaining money and the optimal amount of money held by individuals. If the opportunity cost of bank visits is higher for prime-age people, which seems likely, then prime-age people will hold more money relative to their transactions than will people younger or older. We also test this hypothesis in Section I.[2]

Richard A. Easterlin (1987) and Mark C. Berger (1985) argue that larger cohorts face a lower wage rate, on average, because of the increased competition for jobs. If this is the case, then the size of the cohort should affect the labor-force participation of individuals in the cohort. People in a large cohort will work less if the substitution effect dominates (other things equal) and will work more if the income effect dominates. We examine the relationship between co-

shows that the data are unable to distinguish statistically between a static model including distributional effects and one allowing for dynamics but excluding distributional effects.

The advantage of the age-distribution data used in the present study is that there is considerable variation in the data over the sample period. Thus, there may be a good chance of picking up the effects of the changing age distribution on coefficients in macroeconometric equations.

[2]Thomas Mayor and Lawrence Pearl (1984) estimate, using cross-section census data for the 1950–1970 sample period, that 7.3 percent of the increase in the velocity of money was due to the decrease in the population's median age.

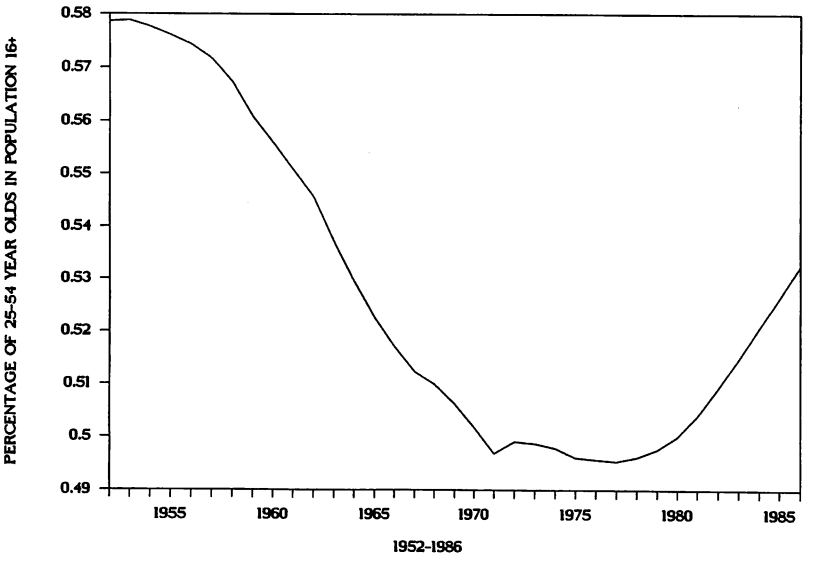

FIGURE 2. PERCENTAGE OF PRIME-AGE PEOPLE IN WORKING-AGE POPULATION

Source: *Current Population Reports*, Series P-25, numbers 310, 519, 917, 965, and 985 (U.S. Department of Commerce, Bureau of the Census).

hort size and labor-force participation in Section II.

I. Consumption, Housing Investment, and Money Demand

A. *The Methodology*

Divide the population into J age-groups. Let $D1_{it}$ be 1 if individual i is in age-group 1 in period t and 0 otherwise; let $D2_{it}$ be 1 if individual i is in age-group 2 in period t and 0 otherwise; and so on through DJ_{it}. Consider the following equation:

$$(1)\quad C_{it} = \mathbf{X}_{it}\boldsymbol{\beta} + \gamma + \alpha_1 D1_{it} + \cdots + \alpha_J DJ_{it} + U_{it}$$

$$i = 1,\ldots,N_t \quad t = 1,\ldots,T$$

where C_{it} is the dependent variable (e.g., consumption or money demand of individual i in period t), $\mathbf{X}_{it}$ is a $1\times k$ vector of explanatory variables excluding the constant, $\boldsymbol{\beta}$ is a $k\times 1$ vector of coefficients, and U_{it} is the error term. The constant term in the equation is $\gamma + \alpha_j$ for an individual in age-group j in period t. N_t is the number of people in the population in period t.

Equation (1) is restrictive, because it assumes that $\boldsymbol{\beta}$ is the same across all individuals, but it is less restrictive than a typical macroeconomic equation, which also assumes that the constant term is the same across individuals. Given $\mathbf{X}_{it}$, C_{it} is allowed to vary across age-groups in equation (1). Because most macroeconomic variables are not disaggregated by age-groups, we are unable to test for age-sensitive $\boldsymbol{\beta}$'s. For example, suppose that one of the variables in $\mathbf{X}_{it}$ is Y_{it}, the income of individual i. If the coefficient of Y_{it} is the same across individuals, say β_1, then $\beta_1 Y_{it}$ enters the equation, and it can be summed in the manner discussed in the next paragraph. If the coefficient differs across age-groups, then the term entering the equation is $\beta_{11}D1_{it}Y_{it} + \cdots + \beta_{1J}DJ_{it}Y_{it}$. The sum of a variable like $D1_{it}Y_{it}$ across individuals is the total income of individuals in age-group 1, for which data are not generally available. One

is thus restricted to assuming that age-group differences are reflected in different constant terms in equation (1).

Let N_{jt} be the number of people in age-group j in period t, let C_t be the sum of C_{it}, let $\mathbf{X}_t$ be the $1 \times k$ vector whose elements are the sums of the corresponding elements in $\mathbf{X}_{it}$, and let U_t be the sum of U_{it}. Given this notation, summing equation (1) yields

$$(2) \quad C_t = \mathbf{X}_t\boldsymbol{\beta} + \gamma N_t + \alpha_1 N_{1t} \cdots + \alpha_J N_{Jt} + U_t \qquad t = 1,\ldots,T.$$

Letting lowercase letters represent per capita terms, equation (2) may be rewritten as:

$$(3) \quad c_t = \mathbf{x}_t\boldsymbol{\beta} + \gamma + \alpha_1 p_{1t} \cdots + \alpha_J p_{Jt} + u_t \qquad t = 1,\ldots,T$$

where, for example, $p_{jt} = N_{jt}/N_t$.

A test of whether age distribution matters is simply a test of whether the α_j coefficients in equation (3) are significantly different from zero.[3] If the coefficients are zero, one is back to a standard macroeconomic equation. Otherwise, given $\mathbf{x}_t$, c_t varies as the age distribution varies. Since the sum of p_{jt} across j is 1 and there is a constant in the equation, a restriction on the α_j coefficients must be imposed for estimation. The age-group coefficients are restricted to sum to zero, $\sum_{j=1}^{J}\alpha_j = 0$; if the distributional variables do not matter, then adding them to the equation will not affect the constant term.

B. *The Data*

The age-distribution data are from the *Current Population Reports*, Series P-25 (U.S. Department of Commerce, Bureau of the Census). The data from the census surveys, which are taken every ten years, are updated yearly using data provided by the National Center for Health Statistics, the Department of Defense, and the Immigration and Naturalization Service. The data are estimates of the total population of the United States, including armed forces overseas, in each of 86 age-groups. Age-group 1 consists of individuals less than one year old, age-group 2 consists of individuals between one and two years of age, and so on through age-group 86, which consists of individuals 85 years old and over. The published data are annual (July 1 of each year). Because the empirical models we estimate are quarterly, we have constructed quarterly population data by linearly interpolating between the yearly points.[4]

We consider 55 age-groups in this study: ages 16,17,...,69, and ≥ 70. The "total" population, N_t, is taken to be the population at least 16 years of age. In terms of the above notation, we have created 55 p_{jt} variables ($j = 1,\ldots,55$), where the 55 variables sum to 1 for a given t.

C. *Constraints on the α_j Coefficients*

It is obviously not sensible to estimate 55 unconstrained α_j coefficients. For the basic results in this paper we have imposed two restrictions on the age-group coefficients. The first, as mentioned above, is that they sum to zero. The second is that they lie on a second-degree polynomial. The second-degree-polynomial constraint allows enough flexibility to see whether the prime-age groups behave differently from the young and old groups while keeping the number of unconstrained coefficients small. A second-degree polynomial where the coefficients sum to zero is determined by two coefficients, and so there are two unconstrained coefficients to estimate per equation. We denote the two variables associated with

[3]Stoker (1986) characterizes this test (that all proportion coefficients are zero) as a test of microeconomic linearity or homogeneity (that all marginal reactions of individual agents are identical). He shows that individual differences or more general behavioral nonlinearities will coincide with the presence of distributional effects in macroeconomic equations.

[4]The quarterly data for the 86 age groups from 1952 through 1988 are available on diskette from the authors upon request.

two unconstrained coefficients, Z_{1t} and Z_{2t}.[5]

We have also examined the robustness of our results to the use of the quadratic restriction. We do this by including four age categories separately in each regression and examining whether the pattern of coefficient values is consistent with the quadratic constraint.

D. *Consumption and Housing Investment*

In this section, we test the hypotheses that age variables have significant explanatory power in consumption and housing-investment equations and that prime-age people consume less relative to their income than those younger or older. To guard against the danger of having the age results depend on a particular model, we estimate equations that are general enough to encompass several different specifications. We start with the equations in the Fair (1984) model and then successively add variables to these equations to make them more general.

The theory behind the specification of the consumption equations in the Fair model is that households choose consumption and labor supply to maximize a multiperiod utility function, possibly subject to a "disequilibrium" constraint regarding the amount that they can work at the current set of wage rates. The explanatory variables include the real value of wealth (A), the after-tax nominal wage (W), the price level (P), the after-tax interest rate (r), the real level of transfer payments (YTR), and a "labor-constraint" variable (Q), which is designed to pick up possible disequilibrium effects.[6] The lagged dependent variable is also included to pick up partial-adjustment and expectational effects.

Labor income is *not* a direct explanatory variable in the consumption equations. If households jointly determine consumption and labor supply and are not constrained in their labor-supply choice, labor income is endogenous. If, on the other hand, households are constrained in their labor supply, then labor supply is no longer a decision variable, and it is appropriate to consider labor income as a determinant of consumption. Q, the labor-constraint variable, is zero when there is full employment (and thus no binding labor constraint on households), and it becomes larger (in absolute value) the more the economy deviates from full employment. In low-employment periods, Q is highly correlated with hours worked, and Q and W together are highly correlated with labor income. Therefore, in low-employment periods the consumption equations reflect the "Keynesian" specification in which income is an explanatory variable. The specification more closely approximates the classical story as the economy approaches full employment.

[5]The age variables enter an equation as $\sum_{j=1}^{55}\alpha_j p_{jt}$, where $\sum_{j=1}^{55}\alpha_j = 0$. The polynomial constraint is $\alpha_j = \gamma_0 + \gamma_1 j + \gamma_2 j^2$, $j = 1,\ldots,55$, where γ_0, γ_1, and γ_2 are coefficients to be determined. The zero-sum constraint on the α_j's implies that

$$\gamma_0 = -\gamma_1(1/55)\sum_{j=1}^{55} j - \gamma_2(1/55)\sum_{j=1}^{55} j^2.$$

The way in which the age variables enter the estimated equation is then

$$\gamma_1 Z_{1t} + \gamma_2 Z_{2t}$$

where

$$Z_{1t} = \sum_{j=1}^{55} j p_{jt} - (1/55)\sum_{j=1}^{55} j \sum_{j=1}^{55} p_{jt}$$

and

$$Z_{2t} = \sum_{j=1}^{55} j^2 p_{jt} - (1/55)\sum_{j=1}^{55} j^2 \sum_{j=1}^{55} p_{jt}.$$

Given the estimates of γ_1 and γ_2, the 55 α_j coefficients can be computed. This technique is simply Shirley Almon's (1965) polynomial-distributed lag technique, where the coefficients that are constrained are the coefficients of the p_{jt} variables ($j = 1,\ldots,55$) rather than coefficients of the lagged values of some variable.

[6]Q is equal to $1 - J^*/J$, where J is the ratio of total hours worked to the total population aged 16 and older and where J^* is an estimate of the full-employment value of J (see Fair [1984] for data references).

We consider four components of household expenditures: service consumption, nondurable consumption, consumer durable expenditures, and housing investment. We estimate five specifications for the service and nondurable consumption equations. The first specification for each equation includes all the variables described previously, plus the age variables; the second specification drops variables whose coefficient estimates are of the "wrong" sign (see Fair, 1990);[7] the third specification adds real disposable income to the variables from the first specification; the fourth specification adds values of all the explanatory variables from the third specification lagged once (except the age variables); and the fifth specification adds values of all the explanatory variables from the third specification led once (except the asset variable, the lagged dependent variable, and the age variables).

The third, fourth, and fifth specifications are clearly more general than the Fair-model specification. Adding disposable personal income in the third specification incorporates the Keynesian specification. Adding the lagged values in the fourth specification incorporates a richer dynamic structure. For example, David F. Hendry et al. (1984) show that adding the lagged dependent variable and lagged values of all the explanatory variables is quite general in that it encompasses many different types of dynamic structures. Adding lead values in the fifth specification allows for the possibility that agents form rational expectations (see Fair, 1990).[8]

All the equations were estimated by two-stage least squares. The first-stage regressors include the main predetermined variables in the Fair model (see table 6-1 in Fair [1984] for the first-stage regressors; see appendix A in that book for the construction of all the variables). The data are quarterly and are seasonally adjusted where appropriate. The estimation period for the equations with lead values was 1954:1–1988:3. For the other equations, the estimation period ended in 1988:4. The significance of the age variables (i.e., the joint significance of Z_{1t} and Z_{2t}) was tested using a chi-square test.[9]

[7]Because of collinearity among explanatory variables and the fairly large number of variables per equation, one would not expect all the coefficients to be estimated precisely. This is an argument for leaving in variables that are not statistically significant at, say, the 5-percent level if their coefficient estimates have the expected sign. Variables with wrong signs were dropped to see if the age results were sensitive to the inclusion of these variables. It will be seen that the results are not.

Note also that the nominal wage (W) and the price (P) have been included separately rather than as W/P. There are two reasons for this. One is to see whether the data support opposite signs on the variables. The second, and more important, reason is that one does not necessarily expect W and P to enter as W/P. The P used for each equation is the own price of the good (the price deflator for services for the services equation, the price deflator for nondurables for the nondurables equation, and so on), and in principle the prices of the other goods should be included in the equation as well. There is, however, too much collinearity among the various price deflators to pick up sensible effects in the aggregate data. Therefore, only the own price was included (but included separately).

[8]When future values for, say, period $t+i$ appear as explanatory variables, Lars Hansen's (1982) method-of-moments estimator can be used. Basically, this estimator is a modification of the two-stage least-squares (2SLS) estimator to account for the moving-average property (of order $i-1$) of the error term. If the lead values are only for period $t+1$ and if equation (A2) in Fumio Hayashi and Christopher Sims (1983) is used to estimate Hansen's "$\mathbf{M}$" matrix, which is commonly done, then Hansen's estimator and 2SLS are the same. For the work in this paper, only $t+1$ was used, and so only the 2SLS estimator was needed.

[9]The chi-square test is as follows. The 2SLS objective function is $\mathbf{u}'Z(Z'Z)^{-1}Z'\mathbf{u} = S$, where $\mathbf{u}$ is a $T \times 1$ vector of error terms and $\mathbf{Z}$ is a $T \times K$ matrix of first-stage regressors; u is a function of the coefficients and the endogenous and predetermined variables in the equation. If the equation is estimated under the assumption of first-order serial correlation of the error term, which is done here for nondurable consumption, u is a nonlinear function of the coefficients if the serial correlation coefficient is counted as a structural coefficient, which is the treatment here. Assume that there are r restrictions on the coefficients. (In the present case, there are two zero restrictions.) Let S^* be the value of S when the restrictions are not imposed, and let S^{**} be the value of S when the restrictions are imposed. Let $\hat{\sigma}^2$ be the estimate of the variance of the error term in the unrestricted case. Then $(S^{**} - S^*)/\hat{\sigma}^2$ is asymptotically distributed as chi-square with r degrees of freedom. A general proof of this is in Donald T. M. Andrews and Fair (1988). If the ordinary least-squares estimator is used, which it is in a few cases in what follows, S is simply $\mathbf{u}'\mathbf{u}$.

TABLE 1—ESTIMATES OF THE SERVICE AND NONDURABLE CONSUMPTION EQUATIONS

	Consumption of services (CS/POP)			Consumption of nondurables (CN/POP)		
Explanatory variable	(1)	(2)	(3)	(1)	(2)	(3)
Constant	0.0773 (3.10)	0.0700 (2.85)	0.0689 (2.61)	0.614 (5.55)	0.606 (5.33)	0.671 (4.40)
LDV	0.845 (22.19)	0.843 (22.10)	0.822 (18.08)	0.199 (1.70)	0.202 (1.67)	0.210 (1.75)
$(A/\text{POP})_{-1}$	0.00064 (1.38)	0.00103 (2.63)	0.00037 (0.69)	0.00335 (4.29)	0.00335 (4.29)	0.00347 (4.24)
r	−0.00279 (5.59)	−0.00253 (5.36)	−0.00262 (5.00)	0.00150 (1.24)	—	0.00175 (1.33)
YTR/POP	0.0500 (1.43)	0.0266 (0.84)	0.0358 (0.95)	0.0800 (1.72)	0.0738 (1.61)	0.116 (1.42)
W	−0.065 (0.52)	0.130 (4.74)	−0.100 (0.77)	0.669 (5.26)	0.599 (5.09)	0.777 (3.31)
P	0.126 (1.59)	—	0.150 (1.81)	−0.468 (4.78)	−0.402 (4.76)	−0.541 (3.27)
Q	0.143 (3.14)	0.117 (2.75)	0.117 (2.20)	0.321 (3.48)	0.353 (3.92)	0.384 (2.60)
YD/POP	—	—	0.020 (0.89)	—	—	−0.037 (0.55)
Z_1	−0.0788 (4.73)	−0.0680 (4.46)	−0.0779 (4.73)	−0.131 (4.78)	−0.141 (5.30)	−0.145 (3.91)
Z_2	0.00157 (4.61)	0.00139 (4.32)	0.00156 (4.60)	0.00256 (4.40)	0.00276 (4.91)	0.00285 (3.62)
$\hat{\rho}$:	—	—	—	0.639 (6.58)	0.649 (6.67)	0.635 (6.47)
SE:	0.00532	0.00534	0.00525	0.00646	0.00641	0.00656
X^2:	24.11	21.30	24.29	34.61	39.75	21.41
α_1:	0.50	0.39	0.49	0.89	0.94	0.96
α_{55}:	1.00	0.93	0.99	1.55	1.69	1.76
j^*:	25	24	25	26	25	25
α_{j^*}:	−0.41	−0.37	−0.41	−0.66	−0.71	−0.74
X^2 (lags):	—	—	27.46	—	—	11.59
X^2 (leads):	—	—	13.47	—	—	8.46

Notes: The estimation technique is two-stage least squares. Sample period = 1954:1–1988:4 (1954:1–1988:3 when leads are used). Numbers in parentheses are t statistics (in absolute value). Statistics reported at bottom of table:

$\hat{\rho}$: estimate of first-order serial correlation coefficient of the error term;

X^2: test of the hypothesis that the coefficients of Z_{1t} and Z_{2t} are zero; critical values ($d.f. = 2$) are 5.99 at the 5-percent level and 9.21 at the 1-percent level;

j^*: value of j for which α_j is at a minimum;

X^2 (lags): X^2 value for the equation with the lagged values of all the explanatory variables added (including the lagged value of the lagged dependent variable) except for Z_{1t} and Z_{2t};

X^2 (leads): X^2 value for the equation with the values of r, YTR/POP, W, P, Q, and YD/POP led one quarter added.

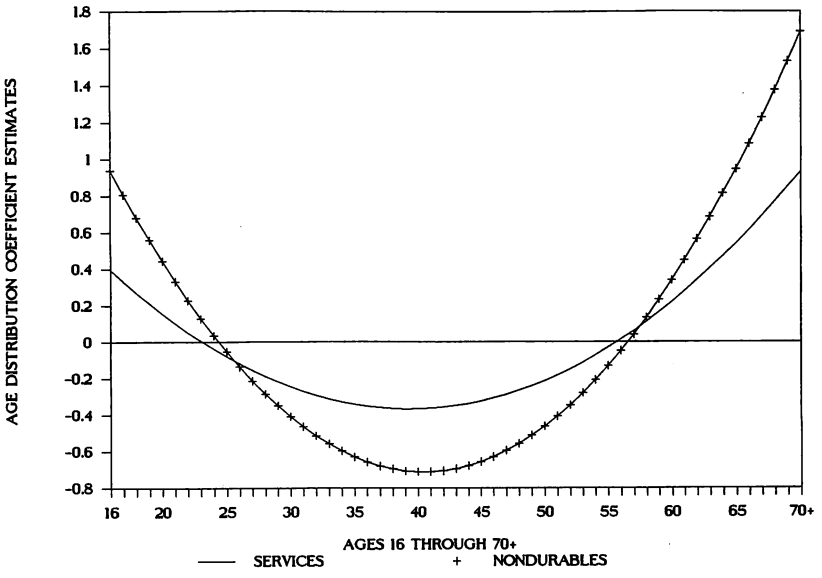

FIGURE 3. CONSUMPTION EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES (SEE TABLE 1)

Table 1 presents estimation results for the service and nondurable consumption equations.[10] Detailed results are presented for the first three specifications, and X^2 values are presented for the fourth and fifth specifications. The results for all specifications strongly support the hypothesis that age variables matter. The lowest X^2 value is 8.46 (significant at the 0.01 level) in the nondurable equation with the leads added. Further, the patterns of the age-variable coefficients are consistent with the life-cycle hypothesis; prime-age people consume less relative to their income than do the young and old. Figure 3 presents the computed α_j coefficients from the second service and nondurable consumption-equation specifications for each of the 55 age-groups. For service consumption the lowest value of α_j occurs at $j = 24$ (age 40), and for nondurable consumption the lowest value occurs at $j = 25$ (age 41).

Going from equations (1) to (2) to (3) in Table 1 has very little effect on the age results. Similarly, adding the lags and leads had little effect on the age results. Since the specification with the lags added is quite general, the age results appear to be quite robust to alternative specifications of the consumption equations.

The equations for expenditures on durables and housing investment in the Fair model are similar to those for service and nondurable consumption, except that they include a lagged stock variable. So long as consumption of housing services is proportional to the stock of housing, the variables that affect consumption, including the age variables, also affect the housing stock. However, unless the actual stock adjusts instantly to the desired stock, an allowance must be made for partial adjustments. Let KH** denote the desired stock of housing if there are no costs of adjusting both the stock and the level of housing investment, where

$$\text{(4)} \qquad \text{KH}^{**} = f(\cdots)$$

[10]The population variable POP in Table 1 that is used to put the equations in per capita terms differs slightly from the population variable N that is used to construct the p_j variables. POP is constructed from monthly data from the Bureau of Labor Statistics, and N is constructed from yearly data from the *Current Population Reports*. N is used in the construction of the p_j variables so that they will sum to 1 across j.

where the arguments in f are the same as for consumption. Two types of partial adjustment are considered. The first is an adjustment to KH** assuming no costs of adjusting investment:

$$(5)\quad KH^* - KH_{-1} = \lambda(KH^{**} - KH_{-1})$$

where KH* denotes the desired stock of housing if there is no cost of adjusting the level of housing investment. The physical depreciation of the housing stock is assumed to be proportional to the size of the stock, with depreciation rate δ. Gross investment in housing (IH) is thus equal to $KH - (1-\delta)KH_{-1}$. Given KH* from (5), desired gross investment is thus

$$(6)\quad IH^* = KH^* - (1-\delta)KH_{-1}.$$

The second type of adjustment is an adjustment of gross investment to its desired value:

$$(7)\quad IH - IH_{-1} = \gamma(IH^* - IH_{-1}).$$

Combining (4)–(7) yields

$$(8)\quad IH = (1-\gamma)IH_{-1} + \gamma(\delta-\lambda)KH_{-1} + \gamma\lambda f(\cdots).$$

This treatment thus adds the lagged dependent variable and the lagged stock of housing to the housing-investment equation, both of which seem to be important explanatory variables in practice.

According to (8), the age variables affect housing investment to the extent that they are arguments in f (i.e., to the extent that they affect housing consumption). One problem with this specification is that the age variables may affect the adjustment parameters λ and γ. In particular, adjustment may be faster for the young than for the old.[11] Unfortunately, as discussed previously, the requisite data are not available to allow parameters other than the constant term to be a function of the age variables, and so it must be assumed that λ and γ do not vary by age. The housing equations may thus be less reliable than those for consumption.

The justification for including the stock of durable goods in the durable-expenditure equation specification is analogous to that described for housing. We estimated four durable-expenditure and five housing-investment equation specifications. (We only estimated four equations for durable expenditures because none of the coefficient estimates in the first equation was of the wrong sign.) Again, the last three specifications are more general than the Fair-model specifications. The values of the explanatory variables in the housing equation are lagged once because housing investment depends in large part on housing-starts decisions made at least a quarter earlier.

Table 2 presents the estimation results for durable expenditures and housing investment. Consider housing investment first. The age variables are highly significant for all the housing-investment equation specifications. Further, the pattern of the α_j coefficients, which are plotted for the second housing equation in Figure 4, is similar to that found for service and nondurable consumption.[12] The lowest value of α_j occurs

[11] Eric A. Hanushek and John M. Quigley (1979) hypothesize that households' consumption of housing in any given period will deviate significantly from their desired level due to the substantial transactions and search costs associated with the housing market. They find (based upon reinterview data gathered on low-income renter households) that young households initially consuming "too little" housing close the gap between actual and desired consumption more rapidly than do older households who are consuming "too much" housing. If there are capital-market imperfections, budget constraints might be less binding for the old than for the young, which would help explain why the young consume too little housing.

[12] Figure 4 suggests that not only do prime-age people consume less housing relative to their income than do those younger and older; but also older people consume more housing relative to their income than do younger people. Louis B. Russell (1982) points out that older people have been major contributors to housing trends because of the rising rate of household (as opposed to individual) incidence in that age-group. The decision not to disband existing households when a spouse retires or dies means that older people continue to occupy the existing housing stock, which requires younger households to buy new rather than existing houses and, thus, for there to be more housing investment.

TABLE 2—ESTIMATES OF THE DURABLE-EXPENDITURE AND HOUSING-INVESTMENT EQUATIONS

	Durable expenditures (CD/POP)		Housing investment (IH/POP)		
Explanatory variable	(1)	(2)	(1)	(2)	(3)
Constant	0.0493	−0.105	0.367	0.399	0.344
	(0.67)	(0.85)	(5.27)	(8.06)	(4.65)
LDV	0.555	0.529	0.874	0.881	0.864
	(6.92)	(6.51)	(31.71)	(32.47)	(29.33)
Lagged stock	−0.0266	−0.0214	−0.0424	−0.0427	−0.0440
	(2.95)	(2.26)	(7.20)	(9.76)	(7.16)
$(A/\text{POP})_{-1}$	0.00270	0.00200	0.00185	0.00127	0.00166
	(3.98)	(2.49)	(3.78)	(3.19)	(3.11)
R	−0.00673	−0.00721	−0.00585	−0.00480	−0.00605
	(5.27)	(5.54)	(6.09)	(6.05)	(6.14)
YTR/POP	0.164	0.090	0.142	0.124	0.119
	(2.69)	(1.17)	(3.57)	(4.17)	(2.55)
W	0.443	0.211	0.154	0.254	0.091
	(2.81)	(0.98)	(2.62)	(7.88)	(1.02)
P	−0.165	−0.002	0.078	—	0.116
	(1.42)	(0.01)	(2.03)		(2.05)
Q	0.247	0.108	−0.023	—	−0.081
	(2.75)	(0.86)	(0.50)		(1.03)
YD/POP	—	0.081	—	—	0.028
		(1.56)			(0.92)
Z_1	−0.0519	−0.0053	−0.157	−0.150	−0.151
	(3.47)	(0.16)	(7.15)	(8.03)	(6.60)
Z_2	0.000706	−0.000291	0.00353	0.00340	0.00344
	(2.29)	(0.41)	(6.60)	(8.12)	(6.31)
SE:	0.00943	0.00934	0.00701	0.00707	0.00702
X^2:	16.83	8.98	59.68	70.57	48.10
α_1:	0.67	0.45	0.60	0.53	0.53
α_{55}:	0.003	−0.72	2.77	2.72	2.75
j^*:	37	55	22	22	22
α_{j^*}:	−0.23	−0.72	−1.00	−0.98	−0.99
X^2 (lags):	—	4.86	—	—	16.42
X^2 (leads):	—	0.77	—	—	33.22

Notes: The estimation technique is two-stage least squares for the durable-expenditure equations and the fifth housing-investment equation. The estimation technique is ordinary least squares for the other housing-investment equations. Sample period = 1954:1–1988:4 (1954:1–1988:3 for durable expenditures when leads are used). Numbers in parentheses are t statistics (in absolute value). For the housing-investment equation, R, YTR/POP, W, P, Q, and YD/POP are lagged one quarter. Statistics reported at the bottom of the table:

X^2: test of the hypothesis that the coefficients of Z_{1t} and Z_{2t} are zero; critical values ($d.f. = 2$) are 5.99 at the 5-percent level and 9.21 at the 1-percent level;

j^*: value of j for which α_j is at a minimum;

X^2(lags): X^2 value for the equation with the lagged values of all the explanatory variables added (including the lagged value of the lagged dependent variable and the lagged value of the lagged stock variable), except for Z_{1t} and Z_{2t};

X^2(leads): X^2 value for the equation with the values of R, YTR/POP, W, P, Q, and YD/POP led one quarter added for durable expenditures and the contemporaneous values added for housing investment.

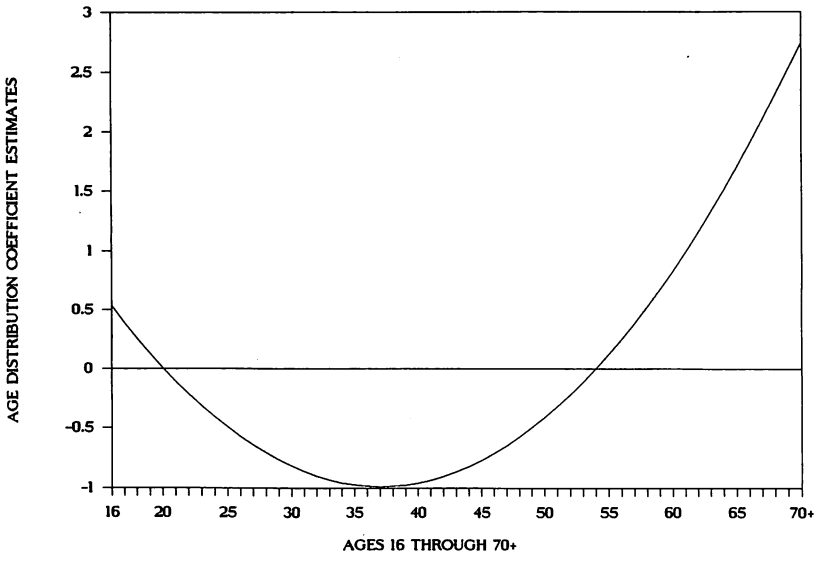

FIGURE 4. HOUSING-INVESTMENT EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES (SEE TABLE 2)

for $j = 22$ (age 38).[13] Again, going from (1) to (2) to (3) has little effect on the age results, as do the additions of the lags and leads.

The results for durable expenditures are mixed. The age variables are significant and have the expected U-shaped pattern for the first equation (plotted in Fig. 5), with the lowest value at $j = 37$ (age 53). However, once we include disposable income and the leads and lags of the other variables, the estimates of the age-variable coefficient are no longer consistent with the life-cycle hypothesis. The age results are thus not robust for durable expenditures.

As noted above, we have examined the robustness of the age results to the use of the quadratic constraint. Instead of imposing the quadratic constraint, the age categories 16–25, 26–35, 36–65, and ≥ 65 were entered separately,[14] with the coefficients on the four variables constrained to sum to zero. This means that three unrestricted age coefficients were estimated rather than two. These results were similar to the results reported above. Consider the equation for each expenditure category with the variables that have the wrong signs dropped. For all four equations, the age variables were significant at the 1-percent level. The chi-square test statistic for the hypothesis that the three age coefficients are jointly zero is 22.13 for service consumption, 39.52 for nondurable consumption, 25.21 for durable consumption, and 64.55 for housing investment.[15] The patterns of the coefficients were also as expected—lower values

[13]N. Gregory Mankiw and David N. Weil (1989), using cross-sectional data for 1970 and 1980, find that the quantity of housing demanded is highest for people between ages 20 and 30 and that it declines after age 40 by about 1 percent per year. Recall that the results presented in this paper suggest that prime-age people consume less housing *relative to their income* than do those older and younger. Therefore, the combined results indicate that prime-age people earn significantly more than those older and younger. This implication is tested directly in the final section of the paper.

[14]The first variable is the percentage of people aged 16–25 in the total population aged ≥ 16; the second variable is the percentage of people aged 26–35 in the total population aged ≥ 16; and so on.

[15]The critical χ^2 value at the 1-percent level with three degrees of freedom is 11.34.

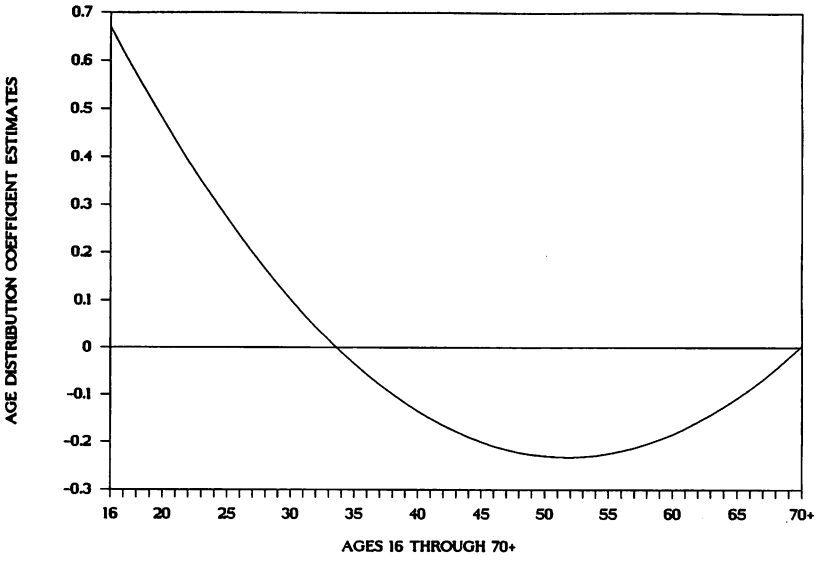

FIGURE 5. DURABLE-EXPENDITURE EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES (SEE TABLE 2)

for coefficients on the second and third age categories than for the first and fourth—except for durable consumption. For durable consumption, the coefficient estimate for the second age category (26–35) was larger than that for the first (16–25), although the other two coefficient estimates were as expected. In general, the results seemed very robust to the different age aggregation.

Overall, the results in Tables 1 and 2 strongly support the hypothesis that age variables matter. Even when lags and leads are added, which are quite general specifications, the age variables retain their explanatory power in all the equations with the exception of durable expenditures. Also, the patterns of the age coefficients are consistent with the life-cycle hypothesis, with the minimum points occurring quite close to where one would expect them. The equation estimates themselves seem reasonably good, with nearly all the coefficients being significant and of the expected sign.

We now turn to money demand, for which one expects the pattern of the age coefficients to be the opposite of that found for the expenditure equations.

E. *Money Demand*

A typical demand-for-money model begins by postulating that the long-run desired level of real money balances (M_t^*/P_t) is a function of real income (Y_t) and a short-term interest rate (r_t). Assume that the equation, in per capita terms, is

$$(9)\quad M_t^*/(P_tN_t) = \alpha + \beta(Y_t/N_t) + \gamma r_t.$$

An adjustment equation is then postulated, in which the adjustment may be either in real terms $[M_t/(P_tN_t)$ adjusting to $M_t^*/(P_tN_t)]$ or in nominal terms $[M_t/N_t$ adjusting to $M_t^*/N_t]$. The results in Fair (1987) strongly support the latter specification, which is used here. The hypothesis is

$$(10)\quad (M_t/N_t) - (M_{t-1}/N_{t-1}) = \lambda(M_t^*/N_t) - (M_{t-1}/N_{t-1}) + \mu_t.$$

Combining (9) and (10) yields

$$(11)\quad M_t/N_t = (1-\lambda)(M_{t-1}/N_{t-1}) + \lambda\alpha P_t + [\lambda\beta(Y_tP_t)/N_t] + \lambda\gamma r_tP_t + \mu_t.$$

If it is assumed that the age variables enter equation (9) as $\gamma_1 Z_{1t} + \gamma_2 Z_{2t}$, then they enter equation (11) as $\lambda\gamma_1 Z_{1t} P_t + \lambda\gamma_2 Z_{2t} P_t$.

Table 3 presents the results of estimating equation (11) with the age variables added. An equation has also been estimated with the lagged values of the explanatory variables added (except for the age variables) to encompass a more general dynamic specification. The age variables in the money-demand regressions are significant at the 10-percent level (although not the 5-percent level) for both specifications. Figure 6 presents a plot of the α_j coefficients for the first equation. The age-variable pattern is as expected. Prime-age people hold more money, other things being equal, than do those younger and older. The largest value of α_j occurs at $j = 23$ (age 39).[16] These estimates can be interpreted as providing at least mild support of the hypothesis that people in their prime working years demand more money relative to their transactions than otherwise because the opportunity cost of their time is higher.

The other coefficient estimates in the money-demand equation seem reasonable. The implied value of λ is 0.167, and the implied value of β is 0.449. The long-run elasticity of money demand with respect to income at the point of means is 0.62.

TABLE 3—ESTIMATES OF THE MONEY-DEMAND EQUATION: DEPENDENT VARIABLE IS MH/POP

Explanatory variable	Coefficient
P	0.253
	(2.77)
MH_{-1}/POP_{-1}	0.833
	(20.06)
$(YD \cdot P)/POP$	0.0749
	(4.29)
$r \cdot P$	−0.00872
	(3.53)
$Z_1 \cdot P$	0.0938
	(2.15)
$Z_2 \cdot P$	−0.00207
	(1.80)
SE:	0.0319
X^2:	4.92
α_1:	−0.38
α_{55}:	−1.59
j^*:	23
α_{j^*}:	0.58
X^2(lags):	4.10

Notes: The estimation technique is two-stage least squares. Sample period = 1954:1–1988:4. Numbers in parentheses are t statistics (in absolute value). Statistics reported at bottom of table:

X^2: test of the hypothesis that the coefficients of Z_1 and Z_2 are zero; critical values ($d.f. = 2$) are 5.99 at the 5-percent level and 9.21 at the 1-percent level;

j^*: value of j for which α_j is at a maximum;

X^2(lags): X^2 value for the equation with MH_{-2}/POP_{-2}, $YD_{-1}P/POP_{-1}$, and $r_{-1}P$ added.

II. Labor-Force-Participation Equations

A typical labor-force-participation equation is of the form

$$(12) \quad L^s_{jt} = X_t\beta + u_t \qquad t = 1,\ldots,T \quad s = \text{m,f.}$$

where L^s_{jt} is the ratio of the number of individuals of sex s in age-group j in the labor force to the total population of people of sex s in age-group j. X_t typically includes the real wage and variables intended to pick up possible "discouraged worker effects."

Although the left-hand-side variables in equations like (12) are disaggregated by age and sex, the right-hand-side variables are typically not age–sex specific. The aggregate real wage is used in place of the more appropriate but unobserved real wage of the particular age–sex group. The implicit assumption in this treatment is that the real wage relevant to age-group j (say W_{jt}) is proportional to the aggregate wage (W_t): $W_{jt} = \lambda_j W_t$. The Easterlin hypothesis suggests that λ_j is a negative function of the percentage of people in age-group j in the total population. Berger (1985) finds that λ_j for baby-boom cohorts is low relative to the λ_j for other age cohorts.

[16] When the four age categories mentioned above were used in place of the quadratic constraint, the age variables were significant at the 5-percent level ($X^2_{[3]} > 10.96$), and the coefficient pattern was as expected.

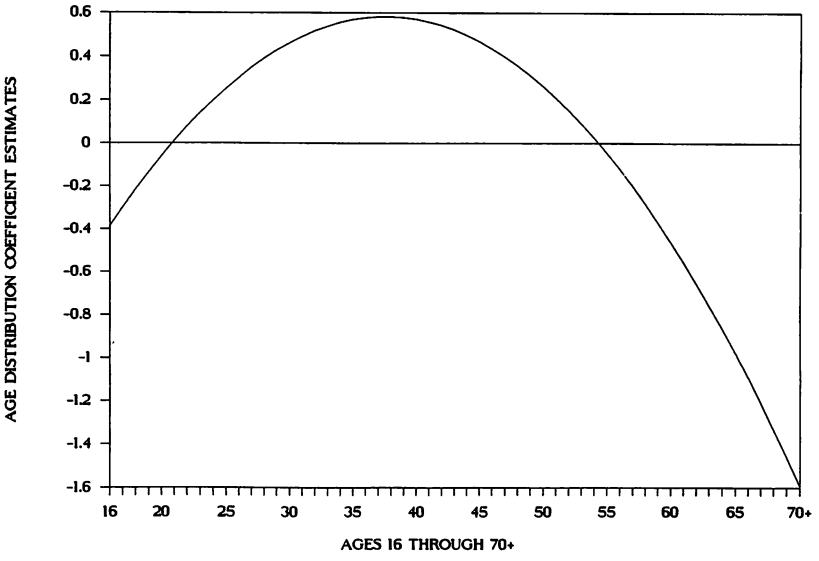

FIGURE 6. MONEY-DEMAND EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES (SEE TABLE 3)

A way to test the Easterlin hypothesis is to postulate that the ratio of the real wage relevant to age-group j (W_{jt}) to the aggregate wage (W_t) is a function of the proportion of people in age-group j (p_{jt}):

$$(13) \qquad W_{jt}/W_t = \gamma_0 + \gamma_1 p_{jt}$$

where γ_1 is negative. Assume that $\beta_1 W_{jt}$ is one of the terms in equation (12). Substituting (13) into (12) results in the terms $\beta_1\gamma_0 W_t$ and $\beta_1\gamma_1 p_{jt} W_t$ in the equation. Since γ_1 is negative, one expects the coefficients of W_t and $p_{jt}W_t$ to be of opposite signs. If the substitution effect dominates, β_1 is positive, and so one expects the coefficient of W_t to be positive and the coefficient of $p_{jt}W_t$ to be negative. The opposite is true if the income effect dominates.

Two aspects of the Easterlin hypothesis should be distinguished. The first is that there is not perfect substitution across age-groups in the labor market, so that more people in an age-group implies a lower average wage for that group. This aspect is represented by equation (13). The second part, termed the relative-income hypothesis, says that young peoples' consumption aspirations are shaped by their parents' living standards. In the face of unfavorable labor-market conditions, a large cohort will adjust demographic and economic behavior in order to maintain its consumption aspirations. More specifically, Easterlin suggests that the baby-boom generation delayed marriage and children and increased labor participation of young women[17] in response to lower average wages.[18] That is, Easterlin postulates

[17]Easterlin notes that, since most young men (in the family-forming ages) are already committed to the labor force, increased labor-force participation will come primarily from young women, although possibly also via moonlighting by the men.

[18]An alternative, sociological, explanation for the increased labor participation (and drop in fertility) of women, as discussed for example in George L. Perry (1977), is based not on the economic incentives brought about by the decline in the relative earnings of the baby-boom generation, but on the changing attitudes about the role of women in society brought about by the "women's movement."

William R. Johnson and Jonathan Skinner (1986) find support for the hypothesis that future divorce probabilities increase current labor supply for married women. They conclude that the rise in the frequency of divorce since 1960 may account for one-third of the unexplained increase in women's postwar labor-force participation.

TABLE 4—ESTIMATES OF THE LABOR-FORCE-PARTICIPATION EQUATIONS

Explanatory variable	Prime-age men L1/POP1	Prime-age women L2/POP1
Constant	0.356	0.0638
	(5.19)	(4.04)
LDV	0.632	0.836
	(8.94)	(19.79)
W	−0.133	0.237
	(4.99)	(4.50)
P	0.0344	−0.0210
	(3.49)	(1.64)
Q	0.0080	0.0366
	(1.07)	(3.21)
Wp_j	0.131	−0.267
	(4.20)	(4.56)
SE:	0.00186	0.00290

Notes: The estimation technique is two-stage least squares. Sample period = 1954:1–1987:1. Numbers in parentheses are t statistics (in absolute value).

that the income effect dominates for women, in which case β_1 should be negative for women, and $\beta_1\gamma_1$ should be positive.

Table 4 presents estimated labor-force-participation equations for prime-age (25–54) men and women. Participation is a function of the after-tax nominal wage (W), the price level (P),[19] the labor-constraint variable (Q), the lagged dependent variable, and the wage rate times p_{jt}. The labor-constraint variable is meant to pick up discouraged-worker effects (discouraged in the disequilibrium sense of not being able to find a job at the current set of wage rates, not discouraged in the sense that the current set of wage rates is low).

For prime-age men, the coefficient estimate for W is negative, and for P it is positive, suggesting that the income effect dominates. The sign of the coefficient on the $p_{jt}W_t$ variable is positive as expected. For prime-age women, the coefficient estimates indicate that the substitution effect dominates. The results thus support Easterlin's hypothesis that relative wages vary inversely with cohort size but fail to support his other hypothesis that the income effect dominates for women.[20]

It is fairly clear from examining the data why the income effect dominates for men. The after-tax real wage generally grew from the beginning of the data set (1952) to about 1974, after which it flattened out. The participation rate of prime-age men fell slightly from 1952 to 1967, fell at a faster rate from 1967 to about 1976, and then flattened out after that. The estimates thus attribute the fall in the participation rate to the rise in the real wage and the flattening out of the participation rate to the flattening out of the real wage. This thus seems to be the income effect at work. The participation rate of prime-age women, on the other hand, has risen fairly steadily over the entire 1952–1987 period, and the estimates are attributing at least some of this rise to the rise in the real wage before 1974. This thus seems to be the substitution effect at work.

III. Caveats

The results in this paper are rather striking. The changing age distribution of the U.S. population has significant explanatory power in consumption, housing-investment, money-demand, and labor-force-participation equations. There seems to be enough variance in the age-distribution data to allow reasonably precise estimates of the ef-

[19]As was the case for the expenditure equations, W and P were entered separately in the equations, rather than as W/P.

[20]Michael L. Wachter (1977) regresses labor-force participation by 14 different age–sex groups on the proportion of the population aged 16–34 in the population at least 16 years of age, the unemployment rate, a time trend, and lagged labor-force participation. He finds that the coefficient estimates for the young-worker variable (aged 16–34) are significantly negative for men aged 25–64 and women aged 45–65+ over the period 1949–1976. However, because the regressions do not include the wage rate, it is not possible to interpret the results in terms of income and substitution effects. Further, it is somewhat unclear as to the expected effects of one age-group's relative size on other age-groups' labor-force participation, which makes the regressions difficult to interpret.

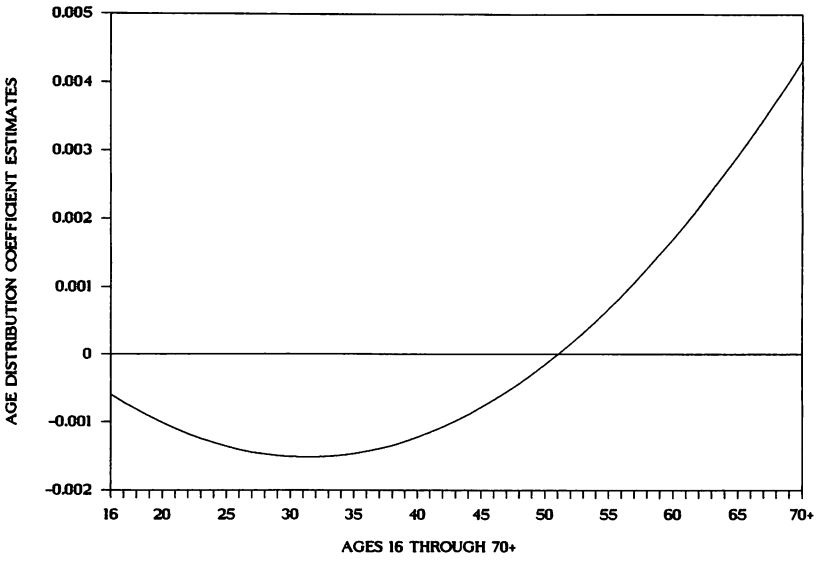

FIGURE 7. REAL-WAGE EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES

Note: Other right-hand-side variables included a constant, time, and the lagged dependent variable.

fects of the age variables on the aggregate variables. The results show that prime-age people consume less relative to their income (including less housing) and demand more money relative to their income than others. The results also show that the labor-force-participation rate of both prime-age men and women is affected by the percentage of prime-age people in the total population. The relationship between cohort size and labor-force participation is positive for men and negative for women.

There are, however, some negative results that should be reported. One might expect that when the age variables are added to an equation explaining real per capita disposable income or the real wage they would be significant and have a sign pattern that implies that prime-age people earn more than those younger and perhaps also more than those older. To test this, Z_{1t} and Z_{2t} were included in a regression of real per capita disposable income on a constant, time, and lagged real per capita disposable income for the 1954:1–1988:4 period. The age variables were not significant ($X^2 = 1.51$). When the two age variables were included in a regression of the real wage on a constant, time, and the lagged real wage, they were significant ($X^2 = 10.91$), but the sign pattern, which is depicted in Figure 7, is not quite as expected. The lowest α_j coefficient occurs at $j = 16$ (age 32), whereas one would expect it to occur at $j = 1$. Also, one would not necessarily expect the curve to keep rising through the last age-group, although this may be because the quadratic constraint is too restrictive.

One might also expect that when the age variables are included in an equation explaining the aggregate personal saving rate they would be significant and have a sign pattern that implies that the saving rate is highest for prime-age people. When Z_{1t} and Z_{2t} were included in a regression of the saving rate on a constant, time, and the lagged saving rate, they were highly significant ($X^2 = 21.03$), but the sign pattern, depicted in Figure 8, is not sensible. The shape of the polynomial implies that those in the

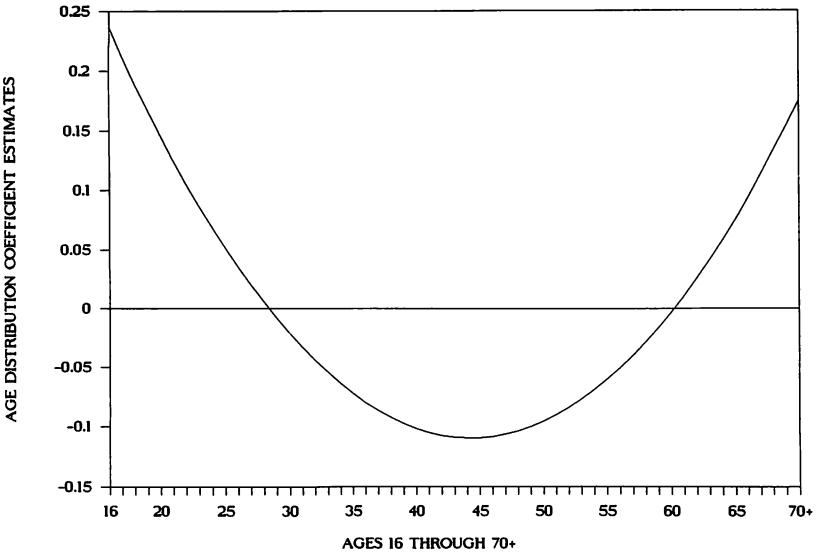

FIGURE 8. PERSONAL-SAVING EQUATION: AGE-DISTRIBUTION COEFFICIENT ESTIMATES

Note: Other right-hand-side variables included a constant, time, and the lagged dependent variable.

prime years save less than those older or younger.

Is there an explanation for these seemingly conflicting results? One possibility is the following. Figure 2 shows that the percentage of prime-age people in the population was low in the 1970's relative to the 1960's and 1980's, which is also true of the state of the economy. There is thus a positive correlation between the percentage of prime-age people in the population and the state of the economy. Since poor states of the economy generally correspond to high saving rates, we can expect there to be a negative correlation between the percentage of prime-age people in the population and the aggregate saving rate. This correlation presumably has nothing to do with long-run life-cycle considerations; it is simply a cyclical fluke in the data.

If this reasoning is valid, then the simple, nonstructural saving-rate regression run above may be spuriously attributing cycle effects to the age variables, whereas the structural equations considered in previous sections may have adequately taken cyclical factors into account with the inclusion of the other variables. Likewise, the disposable-income regression may be contaminated by cycle effects. If so, then the results reported in this section are not of much concern. Nevertheless, the basic results of this paper should be interpreted with some caution. At least part of the significant effects in the previous sections may be from fluke business-cycle correlations rather than from true life-cycle considerations. More work with the age data is clearly needed, but the present results are encouraging.

APPENDIX: VARIABLE DEFINITIONS

A: real value of total net worth of the household sector
CN: real value of the consumption of nondurables
CS: real value of the consumption of services
CD: real value of durable expenditures
IH: real value of housing investment

LDV: lagged dependent variable
L1: total labor force of men 25–54, millions
L2: total labor force of women 25–54, millions
MH: demand deposits and currency of the household sector, billions of current dollars
P: price deflator for CS, CN, CD, or IH in the CS, CN, CD, and IH equations, respectively; price deflator for domestic sales in the MH equation (1982 = 1.0)
POP: population aged ≥ 16 years, millions
POP1: population of men aged 25–54, millions
POP2: population of women aged 25–54, millions
Q: labor-constraint variable
r: after-tax short-term interest rate
R: after-tax long-term interest rate
W: after-tax nominal wage rate
Wp_j: W times the percentage of the population between 25 and 54 years old
YD: real value of disposable personal income
YTR: real value of the level of transfer payments
Z_1: first age variable
Z_2: second age variable

See Fair (1984) for data sources. All real values are in billions of 1982 dollars.

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