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DISSAVING AFTER RETIREMENT: TESTING THE PURE LIFE CYCLE HYPOTHESIS

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ABSTRACT

In this paper, we examine several aspects of saving and dissaving after retirement. First, we argue that existing evidence on bequeathable age-wealth profiles is suspect, and provide new evidence based on longitudinal data indicating that significant dissaving may occur, particularly among single individuals and early retirees. Second, we argue that, in the presence of annuities, estimates of dissaving should be adjusted by including the simple discounted value of benefits in total wealth. Such adjustments reveal relatively little dissaving among any group of retirees. Finally, we test the pure life cycle hypothesis by observing the behavioral response of rates of accumulation to involuntary annuitization, and find empirical refutation of life cycle implications.

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I. Introduction

Does wealth typically decline after retirement? Despite much recent research, this deceptively simple question has remained controversial. Previous investigators seem evenly divided on the issue of whether elderly individuals save or dissave, and no consensus about magnitudes has emerged even among those who agree on the direction of change.

There is as well widespread disagreement about the reasons for asking this question. Some (notably Mirer [1979]) have argued that the life cycle hypothesis is inconsistent with rising or slowly declining wealth after retirement. Others (such as Davies [1981]) have recognized that, in view of uncertainty concerning lifespans, one cannot base a *formal* test of the life cycle hypothesis on this information alone. Such authors have, however, suggested that one could conduct an informal "test" by comparing empirical data with the results of simulations based upon *plausible* parameters values. Finally, one might altogether abandon the hope of inferring motives from information about the age-wealth profile, and instead simply treat such information as valuable *per se*. If, for example, wealth fails to decline rapidly after retirement, intergenerational transfers are likely to be significant. Regardless of motives, this will have strong implications concerning the long run distribution of wealth (see, for example, Loury [1981] and Stiglitz [1978]).

The appropriate definition of "wealth" will depend critically upon which of these purposes one has in mind. Information on bequeathable wealth-age profiles is by itself sufficient for drawing inferences about the magnitude of bequests. However, tests of the life cycle hypothesis must necessarily consider all forms of resources, including annuities (Social Security and pensions). It is therefore somewhat surprising that, with few exceptions (King and Dicks-Mireaux [1982, 1983], Hurd and Shoven [1983]), studies of the age-wealth profile ignore annuities. Nor have any of these authors provided a theoretical discussion of how calculated rates of dissaving should be adjusted in the presence of annuities.

Accordingly, this paper has three objectives. First, we present new evidence on the relationship between age and bequeathable wealth holdings after retirement. While previous studies employ either cross-sectional survey or estate data, our approach is to follow a sample of retired individuals over time. We argue that this methodology is likely to produce superior estimates of dissaving after retirement. We find that bequeathable wealth declines relatively rapidly for single individuals (roughly 3 to 4% per year), while for couples, the evidence is mixed (slight declines, on the order of 1 to 2% per year, are observed for early retirees; otherwise, bequeathable wealth remains relatively constant after retirement). Changes in the composition of bequeathable wealth (specifically, the fraction held as residential housing) are also analyzed.

Our second objective is to develop and implement a technique for calculating meaningful rates of resource depletion when some positive fraction of wealth is held as annuities. Since survival probabilities decline with age, the use of actuarial values (as in King and Dicks-Mireaux [1982, 1983]) builds in a tendency for total wealth to decline quite rapidly after retirement. However, we argue that actuarial discounting is inappropriate for calculating meaningful rates of depletion. Instead, we show that simple discounting of benefit streams is (approximately) appropriate whenever behavior is governed by traditional life cycle concerns. Thus we find, contrary to King and Dicks-Mireaux, that, after adjusting for annuities, neither single individuals nor couples dissave significant fractions of their total resources after retirement.

Of course, this is not a formal test of the life cycle maximization principal. Our third objective is to construct such a test using information on the age-wealth profile. We show that the life cycle model has strong implications about how rates of accumulation and depletion will respond to the imposition of non-discretionary annuities.[1/] Implementation of these tests produces results which are unfavorable to the pure life cycle hypothesis.

The paper is organized as follows. In the next section, we describe the data source which is employed throughout. A discussion of the existing literature on bequeathable wealth-age profiles appears in section 3, along with our new estimates. Theoretical foundations for the valuation of annuity wealth are discussed in section 4, and adjusted estimates of accumulation and depletion are presented. Section 5 describes and implements a test of the life cycle hypothesis based on the behavioral response of changes in wealth to involuntary annuitization. The paper closes with a brief conclusion.

2. The Data

This study employs data from the Longitudinal Retirement History Survey (LRHS), which followed a sample of over 11,000 retirement-aged individuals (58 to 63 in 1968) for a period of ten years, starting in 1969. Some information was also obtained from matching administrative records.

The LRHS collected extensive information on the net worth of respondents. Our measure of bequeathable wealth includes the value of owner occupied housing (net of mortgage liabilities), equity in a business or farm, the net value of other property holdings, cash, and financial assets (including stocks, bonds, bank accounts, checking accounts, and money loaned to other), minus total household debt (excluding mortgage items already counted).[2/]

While extensive in coverage, there is reason to believe that wealth data contained in the LRHS are not of high quality. In general, it is difficult to elicit accurate information about net worth in interview surveys.[3/] A casual inspection of LRHS records indicates substantial misreporting of assets.[4/] Deleting observations for which any component of wealth was incorrectly reported would drastically reduce the sample size, as well as induce a bias of unknown direction. Due to the relative magnitude of housing in the portfolios of most elderly individuals, we did insist that the completion code associated with this item indicated an unambiguous value. This probably biases our sample somewhat towards renters,[5/] although the statistics presented in section 3 suggest that this bias is not large. Throughout the paper, it is important to bear in mind that wealth is poorly reported; we will return to this issue at various points.

Our study also requires extensive information on pensions and annuities. Private and government pension benefits are inferred from income data reported during the sample period. Fortunately, it is possible to distinguish one-shot, lump sum payments from annuities on the basis of recorded responses. For individuals retiring late, benefits from such pensions may commence after 1979 (the youngest respondent is 68 in that year), in which case no income is reported. For such individuals, we supplement income data with survey responses to questions concerning expected levels of future benefits. However, one should bear in mind that private pensions in particular are probably under-reported for late retirees.

Social Security benefits for each year were calculated on the basis of prevailing legislation in that year, using data on covered earnings obtained through matching administrative records. Benefits were calculated on the basis of <u>actual</u> retirement dates for respondent and spouse. For the purpose of this calculation, we assumed that all individuals still working in 1979 retired at the end of that year.

The matching administrative records were also used to calculate a measure of lifetime resources for each respondent. Unfortunately, this information is incomplete, since yearly earnings are only reported up to the taxable maximum. Since the records also indicate the quarter in which the taxable maximum was reached, we were able to extrapolate yearly earnings using the method described by Fox [1976]. The resulting income stream was then accumulated at a 3% rate to a standard age, producing a mesure of lifetime earnings.

Much of our analysis also requires us to know whether a particular individual is retired. Defining retirement is problematic. To reduce contamination arising from the presence of earned income, we created a relatively pristine sample of retirees. Thus, "retirees" report themselves as fully retired in both the retirement year and all successive years, and they report negligible earned income during this period.[6] A retired couple consists of two retired members, while a working couple need only have one worker.

In the following sections, our analysis focuses on the behavior of four samples. To minimize the effects of short run fluctuations, it seemed desirable to look at changes in wealth over relatively long periods. Since the 1973 wave of the LRHS collected very incomplete data on asset holdings, we chose to compare the behavior of retirees and workers over the periods 1969 to 1975, and 1975 to 1979. For the first period, we constructed a sample of households who were retired as of 1969, and deleted all observations which had either dissappeared by 1975 (due to death or attrition), or for which household composition had changed (due to divorce, separation, or death). Similarly, we constructed a sample of households which still included working members as of 1975, and used these as a basis of comparison.[7/] Note that our households are pre-selected on the basis of "survival", and presumably over-represent healthy individuals. This probably biases our estimate of asset decumulation down a bit relative to the correct number for the entire population, but should not affect the comparison of workers and retirees. The second period (1975 to 1979) received identical treatment. Our basic samples consisted of 574 households retired by 1969 (270 single individuals, 504 couples), 1360 households still working in 1975 (240 single individuals, 1120 couples), 1047 households retired by 1975 (173 single individuals, 864 couples), and 497 households still working in 1979 (96 singles, 411 couples).

Finally, all variables have been deflated to 1975 dollars. This, of course, affects the interpretation of magnitudes reported in the following sections.

3. Bequeathable Wealth

Although information about the bequeathable wealth-age profile does not by itself allow us to discuss the plausibility of life cycle motives, it is nevertheless of significant independent interest. In this section, we review the existing literature on dissaving among the elderly, arguing that previous studies suffer from significant biases. New estimates of dissaving from bequeathable wealth are then presented.

A. Previous Studies

Three different types of data sources have been used to estimate the extent of dissaving during retirement. These are: (1) interview surveys of saving among the aged, (2) cross-section interview surveys of net worth, and (3) estate data. We consider these in turn.

Typically, data from interview surveys of saving among the aged (Lydall [1955], Projector [1968], and Mulanaphy [1974]) have found positive, or only slightly negative rates of accumulation. These findings can be criticized on several grounds. First, savings are defined by observable transactions. Thus, all capital gains and losses (including those induced by inflation) are ommitted. Second, the data are highly aggregated. Both Projector and Lydall group all aged individuals (those over 65) together in a single category. Undoubtably, many of these are still working, perhaps saving at a rapid rate in anticipation of retirement. This problem is compounded by the fact that mean values are reported--a small (perhaps wealthy) fraction of the sample saving large amounts may, in such as a calculation, dominate the dissaving of a much larger fraction. Thus, the percentage of retirees dissaving at reasonably rapid rates may be much larger than these numbers would suggest.

A number of investigators, including Lydall [1955], Projector and Weiss [1966], Smith [1975], Mirer [1979], and King and Dicks-Mireaux [1982] have attempted to infer the bequeathable wealth-age profile from cross-section interview surveys of net worth. With the exception of King and Dicks-Mireaux, these studies confirm the findings reported above. However, this approach encounters a variety of difficulties.

First, none of these studies distinguish between workers and retirees. Physical assets understate the total wealth (human and non-human) available to non-retired individuals. Since the proportion of fully retired individuals in a cohort rises with the age of that cohort, this builds in a spurious positive correlation between observed wealth and age.[8/]

To illustrate the potential significance of this affect, we regressed total bequeathable wealth on age and lifetime resources for four subsamples (all single individuals, retired single individuals, all couples, and retired couples), using cross-section data from the 1975 wave of the LRHS. We chose the 1975 wave for two reasons: (1) in 1975, aged of respondent ranges from 64 to 69, which facilitates comparison with other studies,[9/] and (2) in 1969, there was very little spread in age of retirement due to the comparative youth of the sample.[10/] Our results are presented in Table 1.[11/] Point estimates for the entire sample are roughly consistent with previous studies. However, when current workers are excluded, significant dissaving is observed for both single individuals and couples (note, however, that the coefficient is not statistically significant for couples).

Unfortunately, restricting attention to retired individuals within a cross-section induces a sample selection bias. Suppose we know that an individual of age A is retired, but we have not observed his date of retirement. It is straightforward to show that his expected age of retirement is increasing in A.[12/] Thus, all else equal, we would expect older members of a cross-section to have retired later. Differences in age therefore overstate differences in years of retirement (time spent dissaving). This suggests that our estimates understate the true magnitude of dissaving.

A second difficulty encountered by studies employing cross-section interview surveys of net worth is that such surveys implicitly incorporate an important sample selection criterion: only surviving members of a particular cohort are represented. Ex ante, survivors are, on a average, healthier. Thus, as a cohort ages, the survivors will represent an increasingly healthy (in a lifetime sense) fraction of the original sample.[13/] This induces a correlation between age and lifetime health in cross-sections.[14/] Healthier individuals in turn tend to accumulate more wealth to provide for longer retirement periods. As a result, a spurious positive correlation between wealth and age may be observed.

Third, with the exception of King and Dicks-Mireaux, studies employing cross-section surveys of net worth fail to control for lifetime resources. Since wealthier people tend to live longer, older members of any cross-section will, on average, have higher lifetime resources. This problem is compounded by the secular decline in retirement age (older individuals spent more years in the labor force). Rising productivity generates an offsetting "cohort effect"--on average, older members of any cross-section will have worked during periods of lower wages. The net effect is ambiguous; age may be positively or negatively correlated with age in cross-section.[15/]

King and Dicks-Mireaux recognize the importance of controlling for lifetime earnings, and employ the ratio of net worth to "permanent income" as their dependent variable. While an improvement over previous techniques, this fails to correct properly for the first two sources of bias mentioned in the preceeding paragraph. Most obviously, since permanent income is a yearly figure, no adjustment is made for length of working life. In addition, this variable is constructed in a manner which fails to adjust for the correlation between wealth and survival probabilities. Specifically, permanent income is inferred from a cross-section regression explaining current earnings. Since retired individuals have no current earnings, the estimates are driven by the earnings of younger (and therefore, since the cohort effect is corrected for, lifetime poorer) individuals. This builds in a tendency to underpredict the permanent income of elderly individuals, or equivalently to understate the extent of dissaving.

Finally, we consider studies based on estate data. Since Atkinson [1971], Atkinson and Harrison [1978], and Brittain [1978] use this data to generate cross-section estimates of the age-wealth relation, their analyses suffer from the problems described above. In fact, different sample selection criteria imply that, in some cases, the bias will be much worse. For example, information on young individuals is observed only if those individuals die young. Since early death is highly correlated with poor health, there will be a strong correlation between age and lifetime health in such samples. In addition, estate data is heavily truncated, providing no information on a very large number of individuals who die with relatively little net worth. In effect, any individual who dissaves too rapidly is automatically excluded from these samples.

Shorrocks [1975] used a somewhat different approach, estimating the age-wealth relationship from estate data by following a particular cohort over time. While he corrects for potential biases based upon the correlation between wealth and survival probabilities, he does not adjust for the effects of attrition (individuals who dissave sufficiently never show up in estate data), and therefore understates the rate of resource depletion.

While most of these studies have focused on the relationship between total bequeathable wealth and age, some have also investigated changes in portfolio composition among the elderly. One question of particular interest is how the percentage of net worth held as owner occupied housing changes with age. Attempts to infer an answer to this question based upon cross-section data are fraught with the difficulties mentioned above. Portfolio composition may, for example, be related to total lifetime resources, which is correlated with age in cross-sections (see above). It is therefore not surprising that various studies, such as Mirer and King and Dicks-Mireaux, have reached very different conclusions.[16]

B. New Estimates

Since most objections to analyses of cross-section data are based on the premise that individuals at one age are systematically different from individuals at another age, one possible solution is to follow the same individuals over time, observing changes in their net worth. Thus, Mirer concludes that longitudinal data from retirement to death would be "ideal" for determining wealth holding profiles.[17] Diamond and Hausman [1980] have previously employed the National Longitudinal Survey (NLS, or Parnes data) to study individual savings behavior, in part generating an estimate of asset decumulation after retirement. Like the LRHS, the NLS followed a sample of households for a period of ten years; however, NLS respondents are, on average, much younger.[18] Thus, Diamond and Hausman's estimates of decumulation are based on a relatively small,[19] and perhaps atypical[20] sample of retirees. With the completion and availability of the LRHS, it is now possible to supplement the existing literature with new estimates based on more complete longitudinal data for the early retirement period. Our first objective is to provide this evidence.

While the use of panel data does allow us to overcome a variety of difficulities encountered by other approaches, it also raises a new set of problems. First, estimates are very sensitive to macroeconomic events. For example, in a period of supra (sub) normal stock market returns, respondents may experience significant <u>unanticipated</u> accumulation (depletion) of net worth (more on this below). The data, however, provide no way of distinguishing motives. It is worth noting that analyses of cross-section data encounter a similar difficulty, since different cohorts have encountered systematically different patterns of unanticipated gains and losses over the life cycle. Within the current context, we can partially correct for this effect by examining evidence based upon macroeconomically distinct time periods (specifically, we use 1969 through 1975, and 1975 through 1979). In addition, we can, for each period, isolate the net effect of retirement on accumulation by contrasting the behavior of retirees and workers.

A second problem concerns sample selection. For each period, our analysis is confined to households who "survived" the entire period. Presumably, this implies that our data over-represent healthy, wealthy, and domestically stable households. In addition, our requirement that households be retired at the beginning of the period, combined with the relative youth of respondents, implies that the sample is skewed towards early retirees.[21] It is critical to realize, however, that although our sample may be somewhat atypical relative to the entire population,[22] there is no reason to believe that our selection criteria bias estimates of dissaving for this group. The great advantage of panel data is that, by following the same households over time, we can hold exogenous factors (however selected) constant. In contrast, for cross-sections, dissaving is inferred from differences in the net worth of households of different ages, who are implicitly selected according to different criteria. We conclude that panel data, while not perfect, provides a superior source of evidence on asset accumulation.

We begin by inspecting the time pattern of mean bequeathable wealth for each of our subgroups. Results are presented in Table 2.[23] Between 1969 and 1975, net worth declines by 21.1% ($3176) for retired individuals, and 22.8% ($7923) for retired couples. In the later period (1975 to 1979), it declines by 6.8% ($1393) for retired individuals, and rises by 4.1% ($2466) for retired couples. These figures are consistent with a 3 to 4% yearly decline during the first period, and either a 2% yearly decline or 1% yearly rise in the second period. It is difficult to determine whether differences between periods are attributable to sample differences (early vs late retirees), or to changing macroeconomic circumstances.

It is noteworthy that, for each subgroup of working households, net worth always moves in the same direction as it does for the corresponding retired subgroup. In fact, it falls for all groups, except for couples between 1975 and 1979. This by itself is not surprising; hump-shaped income profiles may cause wealth to begin its decline prior to retirement. King and Dicks-Mireaux also find some evidence of dissaving within the pre-retirement group. However, since income falls discontinuously at retirement, the life cycle hypothesis at minimum predicts that the rate of accumulation (depletion) should fall (rise) at that time.[24/] Is this prediction consistent with the data?

For single individuals, there is very little difference in either period between the absolute dollar value dissaved by retirees and workers. However, since early retirees tend to be relatively poor, differences between rates of dissaving are substantial (mean net worth of workers fell 9.4% between 1969 and 1975, and 2.3% between 1975 and 1979). For couples, differences between both rates of change and absolute dollar values dissaved were substantial.[25/] In interpreting these numbers, it is important to recall that the subgroups are based on different sample selection criteria, and differences may therefore reflect heterogenous behavioral propensities.

One puzzling aspect of Table 2 is the precipitous decline between 1969 and 1975 in the net worth of both single individuals and couples still working in 1979. During this period, mean dissaving of households retiring in the more distant future exceeded that of any other groups. This observation seems inconsistent with life-cycle behavior; we will return to it at various point.

For a number of reasons, we are dissatisfied with estimates of accumulation and depletion based on mean values of net worth. Most importantly, these estimates will be heavily influenced by the potentially atypical behavior of households with high initial wealth. Suppose, for example, that the behavior of households i is given by[26]

$$W_{t,i} = \beta_i W_{t-1,i} \quad ,$$

where $W_{t,i}$ is bequeathable wealth in period t. Our estimate of the population dissaving rate, $\hat{\beta}_1$, is:

$$\hat{\beta}_1 = \frac{\bar{W}_t}{\bar{W}_{t-1}} \tag{1}$$

$$= \sum_{i}^{N} \frac{W_{t-1,i}}{N\bar{W}_{t-1}} \beta_i \quad .$$

That is, $\hat{\beta}_1$ is a weighted average of the β_i's, where the largest weights are accorded to individuals with high initial wealth. Such individuals may, for example, be atypically acquisitive, leading to a high estimated value of β.

A related problem concerns measurement error. Suppose that β_i has a common value, β, for all households, so that true wealth $W^*_{t,i}$ evolves according to

(2) $$W^*_{t,i} = \beta W^*_{t-1,i} .$$

Assume as well that wealth is observed with error:

(3) $$W_{t,i} = W^*_{t,i}\ \varepsilon_{t,i}$$

where $E(\varepsilon_{t,i}) = 1$, and $\varepsilon_{t,i}$ is independent of $W^*_{t,i}$ and $\varepsilon_{\tau,j}$ for all $(\tau, j) \neq (t, i)$. Then our estimate $\hat{\beta}_1$ can be written as

$$\hat{\beta}_1 = \sum_{i}^{N} \frac{W_{t-1,i}}{N\bar{W}_{t-1,i}} \tilde{\beta}_i$$

where

$$\tilde{\beta}_i = \beta\ \varepsilon_{t,i}/\varepsilon_{t-1,i} .$$

$\hat{\beta}_1$ is a consistent estimator of β. However, since it is a ratio of stochastic terms, its small sample properties are suspect. In particular, observations with a high value of $\varepsilon_{t-1,i}$ (and therefore a lower value of $\tilde{\beta}_i$) will receive greater weight ($W_{t-1,i}/N\bar{W}_{t-1,i}$ will be higher). We would therefore expect our estimate of β to be biased downwards, towards high dissaving.

These considerations suggest that we should accord equal weight to the dissaving rate of each household. One alternative is to calculate the mean rate, $\hat{\beta}_2$:

$$\hat{\beta}_2 = \frac{1}{N} \sum_i \frac{W_{t-1,i}}{W_{t,i}}$$

(where N is the number of observations). When wealth is observed with error, this technique will produce inconsistent estimates of β. In particular, it is straightforward to verify that, under the appropriate regularity conditions,[27]

$$\text{plim } \hat{\beta}_2 = \beta \, E\left(\frac{1}{\varepsilon_{t,i}}\right) \quad ,$$

which generally exceeds β. The difficulty again arises from the appearance of a stochastic term in the denominator.

We suggest the following procedure. Equation (1) can be written as

$$\log W^*_{t,i}/W^*_{t-1,i} = \log \beta$$

Substituting (2), we see that

$$\log W_{t,i}/W_{t-1,i} = \log \beta + \log \varepsilon_{t-1,i} - \log \varepsilon_{t,i} \quad .$$

If the measurement error terms are, for example, independent[28] and log normal, then the mean observed log rate of accumulation is an unbiased estimator of the log of β. With population heterogeneity, this procedure produces an unbiased estimate of the mean of log β_i, but it is not possible to recover the population mean of β_i itself. However, if the β_i's are reasonably close together (we might expect them to be near unity), the mean of the logs will not be far from the log of the mean.

The problem with the procedure is that it requires us to drop all households for which measured wealth was non-positive in either period t or period $t - 1$. It is important to examine the resulting sample selection bias. If the sample is heterogeneous, the procedure excludes all observations for whom $\beta_i = 0$ or ∞. In addition, if the probability of falsely reporting 0 falls with wealth, then our estimate of the mean of log β_i will be biased upwards.[29]

To determine the potential significance of this effect, we examined the frequency of movements to and from non-positive levels of bequeathable wealth. Our findings are summarized in the second part of Table 2. For most groups (especially couples), the percentage reporting zero wealth was relatively low. Moreover, net movements between positive and non-positive wealth levels are typically quite small (on the order of 1 or 2%), with three exceptions. 6% (net) of retired single individuals moved from positive to nonpositive wealth between 1969 and 1975, and did 15% of retired couples. During the same period, 8% of single individuals who would retire by 1975 moved in the opposite direction. Thus, we observe some tendency for early retirees to completely exhaust their accumulated resources quickly after retirement. We also observe a significant fraction of single individuals accumulating appreciable resources only immeditely prior to retirement.

There is, however, much noise in this data. While net movements between positive and non-positive wealth levels are typically small, the total fraction of households moving in one direction or the other is quite large. To see this, note (in Table 2) that the percentage of households reporting positive resources in two consecutive sample years is substantially smaller than the fraction reporting positive resources in either of those two years alone.

Table 3 presents sample statistics on log W_{75}/W_{69} and log W_{79}/W_{75} for each of our subgroups. Recognizing the conceptual difficulties generated by the sample selection bias described above, we have listed medians, as well as the fraction of each subsample for which bequeathable wealth declines during the period of observation. If inclusion of observations with zero wealth is desired, it is possible to adjust fractile statistics using the percentage movements to and from zero wealth reorted in Table 2.

The results are quite striking, and differ enormously from those based on wealth levels. The mean log rates of accumulation indicate statistically significant dissaving for every retired group, except couples from 1975 to 1979. Positive saving among this group may be an artifact of the precipitous, and probably unanticipated rise in housing prices during the late seventies, combined with relatively widespread home ownership (see statistics below). In contrast, no dissaving is indicated in any currently working group, and in many such cases the estimated saving rates are statistically significant. Note that the "puzzle" of significant dissaving before retirement among late retirees no longer appears. Medians reveal a similar pattern, the only discrepancy in sign arising with respect to single individuals still working in 1979, during the first sample period. Adjustment of medians for movement to and from non-positive wealth would not alter this pattern.

Rates of dissaving for retired single individuals are evidently quite high. Calculated means indicated a yearly decline of between 3 and 6%; medians confirm the lower end of this range. In contrast, couples dissave very little--perhps 1 or 2% per year in the first period (early retirees), and not at all in the second period (although medians indicate that wealth may have risen by as much as 2% per year, the reader should bear in mind the above qualification concerning housing price inflation). The discrepancy between the behavior of single individuals and couples should not be surprising, since couples must provide for the possibility that either member survives for a long time. In addition, it may account for the diversity of previous estimates: Mirer studies couples, while King and Dicks-Mireaux include single individuals.

It is worth noting that saving is observed for a significant fraction (over 40%) of all retired samples, and that dissaving is observed for a significant fraction (over one third) of all non-retired samples. While this phenomenon may reflect heterogeneity of behavior, we are inclined to attribute it primarily to the apparent extent of measurement error.

Only our highest estimates of depletion rates are roughly consistent with the 6.8% figure obtained by Diamond and Hausman. We attribute the magnitude of their estimate to the unrepresentative characteristics of their sample. As mentioned earlier, NLS households are, on average, substantially younger than LRHS households. Individuals retiring during the NLS sample period will, by and large, be early retirees; our results indicate that early retirees tend to over-represent single individuals,[30] and we have seen that single individuals deplete resources more rapidly than couples. In light of our findings, it would seem unwise to conclude on the basis of their study that typical married retirees dissave significant portions of their wealth.

We now examine the evolution of portfolio composition after retirement. Table 4 decomposes bequeathable wealth into four categories: owner occupied housing, business and property, financial assets, and debt (other than mortgages). The last of these categories is insignificant. The extent of homeownership (fraction owner-occupants) is also indicated.

For both single individuals and couples retired by 1969, there is a decline in every significant asset category <u>except</u> housing. The data indicate a slight increase in homeownership for retired individuals during this period, and a slight decline for retired couples.

The behavior of households which were retired by 1975 is more interesting. More or less simultaneously with retirement (1969 to

1975), both single individuals and couples liquidated large amounts of business and property wealth. At the same time, holdings of financial assets rose slightly, while large gains in housing wealth (especially in frequency of home ownership) were registered. This raises the possibility that households liquidated business and property holdings to finance purchases of homes.[31/] During the post-retirement period, there is a slight dip in homeownership for both groups. Evidently, while many households purchase homes at retirement, a smaller but significant number sells homes within a few years subsequent to retirement.

The evidence also appears to indicate that a reasonably stable (perhaps slightly increasing) fraction of bequeathable wealth is held as owner-occupied housing during retirement. This confirms the finding of King and Dicks-Mireaux, contradicting that of Mirer. However, we should emphasize that these data only concern the early retirement period.

4. Annuities

A very large fraction of the total resources available to many retired individuals is locked into annuities (government and private pensions, Social Security). Studies which ignore this important component of wealth fail to provide sufficient information for judging the plausibility of life cycle motives.

It has frequently been argued that the inclusion of annuities would vindicate the hump-shaped wealth-age profile, since the actuarial value of survival contingent claims falls with age (single year survival probabilities decline). Thus, Mirer [1979] concedes that, "to some extent, perhaps a great one for many people, pension and Social Security programs tend to institutionalize the tenets of the life cycle theory." Likewise, King and Dicks-Mireaux [1983] find evidence of "a clear life-cycle pattern" when the actuarial value of annuity claims are included in measures of net worth.

In this section, we argue that actuarial valuation is inappropriate if one wishes to infer an age-wealth profile in order to judge the plausibility of life cycle motives. Elsewhere (Bernheim [1984b]), we have shown that the simple discounted value of future benefits (ignoring the possibility of death) is ordinarily a good approximation to the value (in terms of compensating variation) of an annuity. Here, we establish that simple discounting is also appropriate within the current context. Since this measure changes very little with age, our analysis reverses the conclusions of King and Dicks-Mireaux: the inclusion of annuities reinforces earlier findings that resources decline only slightly, if at all, after retirement.

A. Theoretical Considerations

Actuarial valuation of annuities is appropriate under either of two conditions: (1) households are risk neutral, or (2) households have access to competitive annuity markets. The first of these conditions is unreasonably restrictive, and generates absurd behavioral predictions.[32] Under the second condition, there is a very simple test of pure life cycle motives: do households hold positive levels of bequeathable wealth at all? In fact, if annuities yeld any return in excess of the interest rate, pure life cycle consumers will annuitize 100% of their resources,[33] and the notion of dissaving will be vaccuous. Thus, if we wish to use evidence on rates of dissaving to test the pure life cycle hypothesis, we must assume a complete absence of annuity markets.[34]

Under the assumptions of missing annuity markets and risk aversion, the value of an annuity will exceed its actuarial value by a risk premium. Our current task is to determine what this observation implies about the appropriate computation of age-wealth profiles.

We will assume the constant elasticity, intertemporally separable form of lifetime utility,

$$\frac{1}{\alpha} \int_0^\infty e^{-\lambda t} C_t^\alpha \, dt \quad , \tag{3}$$

where λ captures the effects of discounting both through the pure rate of time preference and survival probabilities.[35] At time 0, the individual is endowed with some level of bequeathable wealth W_0, and receives some annuity payment A_0. Annuity payments grow geometrically at the rate g; the interest rate is r. Thus, the individuals choice is constrained as follows:

$$\int_0^\infty (C_t - A_t)e^{-rt} \, dt \leqq W_0 \tag{4}$$

and

$$W_t = e^{rt}W_0 - \int_0^t (C_\tau - A_\tau)e^{r(t-\tau)} \, d\tau \geqq 0. \tag{5}$$

Ignoring constraint (5) and maximizing (3) subject to (4), we obtain the following first order conditions:

(6) $$C_t = e^{\gamma t} C_0$$

where $\gamma \equiv (r - \lambda)/(1 - \alpha) < r$. 36/ Suppose $\gamma \geqq g$. Then, continuing to ignore (5), it is easy to see that the optimal program is given by (6), along with

(7) $$C_t = (r - g)(W_t + \frac{A_t}{r - g})$$

and

(8) $$(W_t + \frac{A_t}{r - g}) = (W_0 + \frac{A_0}{r - g})e^{\gamma t}.$$

Since this program never violates (5), it is optimal.

The interpretation of (7) and (8) is straightforward: consumption in each period is a constant fraction of total wealth, and total wealth grows at the geometric rate γ. Note, however, that the annuity wealth term, $A_t/(r - g)$, is equal to the simple discounted value of future benefits (ignoring death). Thus, to make inferences about γ (the life cycle parameter of interest) from data on the age-wealth profiles, we should define total wealth to include the <u>simple</u> discounted value of annuities, <u>not</u> the actuarial value. Intuitively, unless an individual plans to consume his principal at some point in the future, he will be indifferent an annuity paying \$1 per year, and an asset worth \$1/r (both generate the same survival contingent income stream).

If $\gamma < g$, the problem is more complex. Ignoring (5), one again obtains (7) and (8), but in this case (5) will be violated for t

sufficiently large (the individual will wish to borrow on future annuity benefits). Along the true optimal program, consumption will obey the first order condition (6) as long as wealth is positive; however, once (5) binds, we will simply have $C_t = A_t$. Let T denote the age at which (5) first binds. Then the first order conditions imply that

$$(9) \qquad \begin{cases} C_t = e^{\gamma t} C_0 & t < T \\ \\ C_t = A_t & t \geqq T \end{cases}$$

From the resource constraint, we have

$$(10) \qquad W_0 = \int_0^T (C_t - A_t)e^{-rt}\, dt.$$

Finally, it is easy to see that despite the binding constraint, consumption must be continuous in time, so that

$$(11) \qquad e^{\gamma T} C_0 = e^{gT} A_0.$$

Equations (9), (10), and (11) together determine C_0 and T, from which the optimal program can be constructed.

In Bernheim [1984b], we calculated the compensating variation associated with the marginal annuity for the case of $\gamma < g$ (using equations (9) through (11)),

$$\left.\frac{dW_0}{dA_0}\right|_{U=U*} = -\frac{1}{r-g}\,[1 - \phi]$$

where

$$\phi = \left[1 - \frac{r - g}{\lambda - \alpha g}\right] e^{(g-r)T} \quad ,$$

and established that $0 \leq \phi < 1$. Intuitively, since (5) may bind at some point, the annuity is worth less than an asset which yields the same yearly survival contingent income. As T goes to infinity (or γ to g), this event becomes more remote, so naturally the value of annuitization approaches $A_0/(r - g)$.

Hypothetical values of the proportional adjustment factor (ϕ) are given in Bernheim [1984b]. For completeness, we reproduce two sample calculations here. We assume that $r = 0.03$, $g = 0$, $\alpha = 0$ (the logarithmic case), and $A_0/(r - g)W_0 = 2$ (i.e., two thirds of total resources are held as annuities).[37] Since λ depends on the rate at which individuals discount future utility, it is the most difficult parameter to gauge. We employ values of 0.05 and 0.07.[38] The formula for γ is given above. Substituting (9) into (10), one finds that T is given by the implicit solution to

$$e^{(g-\gamma)T}(1 - e^{(\gamma-r)T}) \frac{r - g}{r - \gamma} - (1 - e^{(g-r)T}) = \frac{W_0(r - g)}{A_0} \quad .$$

Calculated values of γ, T, and ϕ are pesented in Table 5. Ignoring non-negativity constraints, wealth would decline by 2 and 4% per year, for λ equal to 0.05 and 0.07, respectively. The associated unconstrained intervals are 42 and 27 years. The marginal annuity is worth 89%, and 75% of its simple discounted value, respectively.

Employing a "triangle approximation" for the value of inframarginal units, we find that the associated compensating variations for all annuity holdings are 94%, and 87% of their simple discounted values. In contrast, for these parameter values the actuarial discounted value of a benefit stream is only 37.5% of its simple discounted value.[39]

There is, of course, no reason to believe that it is appropriate to use the compensating variation as a measure of annuity valuation when calculating wealth trajectories (except in the limiting case where the non-negativity constraints never bind). For this reason, we pose the question somewhat differently. Suppose we employ simple valuation; i.e., define total resources,

$$R_t \equiv W_t + A_t/(r - g) \quad ,$$

and calculate rates of dissaving from R_t/R_0 (i.e., pretend the non-negativity constraints never bind). How well will our estimated dissaving parameter,

$$\gamma^r \equiv t^{-1}\ell n(R_t/R_0) \quad ,$$

approximate the parameter of interest (γ)?

Using our characterization of the optimal (constrained) program, it is possible to calculate that

$$W_t = e^{rt}[W_0 - A_0\{e^{(g-\gamma)T}(1 - e^{(\gamma-r)t})/(r - \gamma)$$

$$- (1 - e^{(g-r)t})/(r - g)\}] \quad .$$

Substituting this into the expression for R_t, one can show (after some tedious manipulations) that

$$\frac{R_t}{R_0} = e^{\gamma t}[1 + \psi] \tag{12}$$

where

$$\psi = e^{(r-\gamma)t}\left[\frac{1 - e^{(\gamma-r)t}}{1 - e^{(\gamma-r)T}}\right]\left[\frac{A_0/(r - g)}{R_0} e^{(g-r)T} - e^{(\gamma-r)T}\right] \tag{13}$$

Table 5 presents values of ψ and γ^r calculated for our sets of hypothetical parameter values (where $t = 6$). When $\lambda = 0.05$, ψ is 0.027, which indicates that γ^r understates the "true" rate of dissaving by approximately 1/2% per year. Thus, rather than observing a decline of 2% per year, we should observe "total wealth" falling by 1 1/2% per year. When $\lambda = 0.07$, $\psi = 0.090$, which indicates that γ^r understates the true rate of dissaving by 1 1/2% per year. Thus, "total wealth" would fall by 2 1/2%, rather than by 4% per year.

These calculations suggest that γ^r will, for $\gamma < g$, understate the rate of dissaving, γ. We now prove that this inequality always holds.

Proposition 1: For $\gamma \geqq g$, $\gamma^r = \gamma$. For $\gamma < g$, $\gamma^r > \gamma$.

Proof: The first statement follows trivially from equation (8). We prove the second claim by showing that $\psi > 0$. Straightforward calculations reveal that, for $\gamma < g$, $dC_0/dA_0|_{R_0} < 0$ [40] (intuitively, annuities have a negative income effect since the non-negativity constraint binds; consumption is therefore depressed). Thus, $R_t > R_0 e^{\gamma t}$ (since the right hand side indicates remaining resources in period t if non-negativity constraints are ignored). Taking $t = T$ and rearranging, we see that $A_0 e^{gT}/(r - g) > R_0 e^{\gamma T}$. From equation (13), this is easily seen to imply that ψ is positive.//

Given this result, one possible approach is to adjust γ^r given an assumed value of ψ, corresponding to some set of reasonable parameter values. Unfortunately, ψ depends on γ, so we cannot estimate γ from γ^r without knowing γ itself. Another alternative is to obtain a lower bound on γ, in addition to this upper bound.

How might we obtain a lower bound? One suggestion is to calculate rates of dissaving from W_t/W_0 (as in the preceeding section):

$$\gamma^w \equiv t^{-1} \ln(W_t/W_0) \quad .$$

To motivate this suggestion, ignore (for the moment) non-negativity constraints (equation (5)). Equation (8) will then describe the evolution of total wealth. Simple manipulations reveal that

$$\text{(14)} \qquad \frac{\dot{W}_t}{W_t} = \gamma + (\gamma - g) \frac{A_t/(r - g)}{W_t} \quad .$$

Equation (14) has an important interpretation. If the individual holds no annuities, his bequeathable wealth grows at exactly the rate γ. Supposing as before that $\gamma < g$, as annuities increase, the rate at which bequeathable wealth declines will accelerate.[41] The reason is straightforward: annuity wealth $(A_t/(r - g))$ declines at the rage g; to preserve a total rate of decline of γ, bequeathable wealth must fall at an accelerated rate. Thus, as long as $\gamma < g$, γ^{w} will <u>overstate</u> the extent of dissaving. Note that this is completely contrary to the assertions of earlier authors, who had argued that W_t/W_0 would understate dissaving due to the actuarial decline in annuity wealth.

Of course, the preceeding analysis ignores the non-negativity constraints. It is important to verify that our lower bound on γ is valid even when these constraints are considered explicitly. In particular, we prove:

<u>Proposition 2</u>: When $\gamma < g$, $d(\dot{W}_t/W_t)/dA_0 < 0$.

<u>Proof</u>: Using the accounting identity

$$\frac{\dot{W}_t}{W_t} = r + \frac{A_t - C_t}{W_t}$$

we see that

$$W_t^2 \frac{d(\dot{W}_t/W_t)}{dA_0} = W_t\left(\frac{dA_t}{dA_0} - \frac{dC_t}{dA_0}\right) - (A_t - C_t)\frac{dW_t}{dA_0} \cdot$$

Appropriate substitution from equations (7) through (10) reveals that this is[42/]

$$= A_t\, e^{(g-r)(T-t)}$$

$$\cdot \left[\frac{e^{(r-\gamma)(T-t)} - 1}{r - \gamma}\right]\left[e^{rt} + \left(1 - \frac{r - \gamma}{r - g}\delta\right)\right]$$

$$+ \left[\frac{e^{(r-g)(T-t)} - 1}{r - \gamma}\right]\left[e^{\gamma t}\left(\frac{r - \gamma}{r - g} - 1\right)\delta - e^{rt}(1 - \delta)\right]$$

where

$$\delta = \frac{1 - e^{(g-r)T}}{1 - e^{(\gamma-r)T}} \quad .$$

Using the fact that

$$\frac{e^{(r-\gamma)(T-t)} - 1}{r - \gamma} > \frac{e^{(r-g)(T-t)} - 1}{r - g} > 0 \quad ,$$

it is then possible to show that[43/]

$$\frac{d(\dot{W}_t/W_t)}{dA_0} < 0 \quad ,$$

which is the desired result.//

Of course, if $A_0 = 0$, $\dot{W}_t/W_t = \gamma$, so for $\gamma > g$, $A_0 > 0$ implies $\gamma^w < \gamma$. It is convenient to summarize this conclusion, as well as much of the preceeding analysis, in the following proposition.

<u>Proposition 3</u>:

(i) If $\gamma = g$ or $A_0 = 0$, $\gamma^r = \gamma^w = \gamma$.

(ii) If $\gamma > g$ and $A_0 > 0$, $\gamma^r = \gamma < \gamma^w$.

(iii) If $\gamma < g$ and $A_0 > 0$, $\gamma^w < \gamma < \gamma^r$.

Case (iii) is the most interesting, since (for $g = 0$) it concerns a dissaver who holds positive annuities. For such an individual, depletion of bequeathable wealth will overstate dissaving, while depletion of total wealth (including the simple discounted value of annuity benefits) will understate it.

Which of our two measure, γ^r or γ^w, will be closer to γ? In general, the answer depends upon particular parameter values. We can obtain some feel for magnitudes by using (12), along with the definition of R_t to obtain

$$\frac{W_t}{W_0} = \frac{e^{\gamma t}[1 + \psi] - \xi e^{gt}}{1 - \xi} \tag{15}$$

where

$$\xi = \frac{A_0/(r - g)}{R_0} \quad .$$

Suppose $g = 0$. What happens as ξ rises? Ignoring the effect on ψ, we see that W_t/W_0 falls; in fact, it is equal to zero when $\xi = e^{\gamma t}[1 + \psi] < 1$. Thus, we would expect γ^w to significantly understate γ when the degree of annuitization is high.

The data presented below indicate that ξ is quite high--roughly on the order of 2/3 (while others have found much lower levels of annuitization relative to bequeathable wealth, this is due to the use of actuarial valuation). It is therefore not very surprising that γ^r significantly outperforms γ^w for our hypothetical parameter values. In Table 5, we calculate values of γ^w, using equation (15). Increasing annuitization from zero to two-thirds of total resources accelerates the rate of bequeathable wealth depletion from 2% to 5.2% per year for $\lambda = 0.05$, and from 4 % to 9.3% for $\lambda = 0.07$. In both cases, the true value of γ is much closer to our upper bound, γ^r. By incorporating date on annuuities, we might therefore hope to learn much more about the implied behavioral rate of dissaving.

B. Analysis of the Data

In implementing the ideas described above, we encounter two conceptual difficulties. The first concerns expectations about future annuity benefits. In particular, substantial changes in Social Security legislation took place during the sample period. Should we assume that these were properly anticipated? If we assume myopic expectations at each point in time (constant real benefits from that point forward), Social Security wealth will be quite volatile. However, since by assumption this volatility is unanticipated, resulting changes in wealth should not be counted as saving or dissaving. In such a world, planned dissaving from Social Security is necessarily zero by definition.

In practice, we assume that all changes in Social Security legislation during the sample period were correctly anticipated, and that constant real benefits were expected after 1979. This tends to minimize changes in Social Security wealth induced by legislative action. We also assume that government and private pensions were expected to provide constant real and nominal benefits, respectively.[44/]

A second difficulty concerns the proper treatment of couples. The model described above is out of its depth when household members can die at distinct points in time. If, however, annuities have full assumption of benefits by a surviving spouse, then our conclusion is essentially unchanged: if the household has no bequest motive, and if its members would never want to consume the principal of an asset, then it must be indifferent between that asset and an annuity which pays the same income stream. Thus, simple discounting is still appropriate. If the desire to consume the principal will arise only far in the future, then simple discounting must be a good approximation.

For government and private pensions, we assume full transfer of benefits, so the difficulty dissappears. However, we know that this is counterfactual in the case of Social Security. We resolve this dillemma by decomposing Social Security into two streams: a certain stream (equal to the minimum benefit under any survival contingency), and a contingent stream (equal to the residual). By the preceeding argument, simple discounting is approximately appropriate for the certain stream. In the following analysis, we simply ignore the contingent stream. We suspect that the insurance value associated with this contingent stream does not change enough over time to alter any of our qualitative conclusions.

In Table 6, we present calculation of annuity wealth for the samples described in section 2. The presentation of these numbers is designed to facilitate comparison with the results on bequeathable wealth.

Note that between 1969 and 1975, annuity wealth rises steeply for most pre-retirement groups. Since pensions pay little or no income to such individuals during this period, pension assets effectively earn interest as the date of benefit eligibility approaches (the rise in pension wealth is due solely to this effect; in these calculations, continuing to work does not per se contribute to the value of benefits). Note that this effect is not very significant for working households between 1975 and 1979; evidently, most of these households began to receive benefits prior to full retirement.

For retired groups, annuity wealth changes very little, as expected. During the sample period there are two countervailing effects: legislation increases the real value of Social Security, while inflation erodes the value of private pensions. The first effect is not as large as one might expect, since we assume that future legislative changes are correctly anticipated. Thus, the Social Security wealth stream is relatively flat. Since private pensions are discounted at a much higher rate, Social Security dominates these calculations. Nevertheless, the erosion of private pension values contributes to a slight decline in total annuity wealth.

In Table 7, we combine data on bequeathable wealth and annuities. Due to the size of annuities relative to bequeathable asset, the total wealth-age profile is relatively flat. For retired single individuals, total wealth appears to decline by at most 1% per year. In fact, between 1969 and 1975, total wealth increased for more than half of these households. Retired couples exhibit a slight decline (1 to 1 1/2% annual) in total wealth during the early sample period, but show virtually no change during the later period. In contrast, working households show slight increases (0 to 2%) in total wealth for almost every period and subsample. Note that the "puzzle" concerning the precipitous decline between 1969 and 1975 in the bequeathable wealth of late retirees now acquires a new interpretation: this dissaving simply offset the implicit saving accompanying the approach of pension eligibility.

Contrary to King and Dicks-Mireaux, we have found that evidence of rapid dissaving among the elderly dissapears when annuities are considered. Our calculations based on hypothetical parameter values in a simple life cycle model (Table 5) suggest that the data on bequeathable and total wealth profiles (Table 3 and 7) together are consistent with a behavioral dissaving rate of less than 2% per year.[45] However, as noted before, this does not constitute a formal test of the life cycle hypothesis. In the next section, we investigate the possibility of basing a formal test on information about the age-wealth profile.

5. Testing the Pure Life Cycle Hypothesis

While rates of dissaving may not, by themselves, confirm or refute the life cycle hypothesis, the observed response of these rates to involuntary annuitization may provide a basis for doing so. This suggestion motivates the following analysis.

Returning to our formal model, let us assume that, as an approximation, we can ignore the effect of non-negativity constraints (equation (5)). Equation (14) will than describe the evolution of bequeathable wealth. It is useful to rewrite this as

$$(16) \qquad \frac{\dot{W}_t}{W_t} = \gamma + \xi(g)\,\frac{A_t}{W_t} \quad .$$

where

$$\xi(g) = \frac{\gamma - g}{r - g} \quad .$$

Notice first that the sign of $\xi(g)$ is the same as that of $\gamma - g$. This simply reflects the phenomenon noted earlier: annuitization will accelerate (decelerate) the growth of bequeathable wealth if and only if $\gamma > g$ ($\gamma < g$). We illustrate this pattern in Figure 1. Suppose that two individuals have different behavioral dissaving parameters (γ_1 and γ_2), but that their annuity benefit profiles have a common growth rate, g. If $\gamma_1 > g > \gamma_2$, annuitization will accelerate bequeathable wealth accumulation for individual 1, and slow it for individual 2. Proposition 2 confirms that explicit consideration of the non-negativity constraints does not change this conclusion.

A test based on the behavioral response of accumulation rates to involuntary annuitization should have substantial power against major alternatives. The existence of an operative bequest motive would, for example, imply that annuitization <u>always</u> causes bequeathable wealth to accumulate more rapidly (decline more slowly).[46] A similar implication is generated by more simple minded models, in which households save some constant fraction of current income.

Next, observe that, to a first order aproximation (expanding ξ around $g = 0$),

$$(17) \qquad \xi(g) \approx \frac{\gamma}{r} + \frac{\gamma - r}{r^2} g.$$

The transversality condition guarantees that the coefficient of g is unambiguously negative (in fact, for all $g < r$, $\xi'(g) < 0$; ξ falls as the growth rate of annuity benefits rises). Intuitively, increasing the value of g may shift an individual from the regime in which annuitization accelerates the growth of bequeathable wealth $(\gamma > g)$ to the regime in which the effect of annuitization is reversed $(\gamma < g)$. This is illustrated in figure 1: for $g' > \gamma_1 > g$, individual 1 belongs to the class of consumers who respond to annuitization by accumulating wealth at a slower rate (dashed lines indicate behavioral responses associated with an annuity benefit growth rate of g'). This implication is, as well, presumably testable.

Our data on bequeathable wealth profiles, of course, only allow us to measure discrete changes, rather than continuous rates of change. In moving to our empirical implementation, we must therefore begin by converting (16) into its discrete analog:

$$\frac{W_{t+1}}{W_t} \approx (1 + \gamma) + \xi(g)\,\frac{A_t}{W_t}.$$

For reasons discussed in section 3, we prefer to use the log rate of accumulation as our dependent variable. Since the rate is presumed close to unity for most observations, we can employ the following approximation:

$$\ell n\; W_{t+1}/W_t \approx \gamma + \xi(g)\,\frac{A_t}{W_t}.$$

Finally, using our first order approximation of $\xi(g)$ (equation (17)) and adding a stochastic error term (representing among other things, the effects of the preceeding approximations), we produce our basic specification:

$$\ln W_{t+1}/W_t = \gamma + \frac{\gamma}{r}\frac{A_t}{W_t} + \left(\frac{\gamma - r}{r^2}\right)g\frac{A_t}{W_t} + \varepsilon_t. \tag{18}$$

Given cross-sectional data on bequeathable wealth and annuities (including the growth rate of benefits), one could estimate equation (18), alternatively ignoring and imposing (through a NLLS procedure) the implicit constraints on the coefficients. The model could then be tested by evaluating (statistically) the plausibility of these constraints, and by examining the sign of $\gamma - r$ in the constrained version. We eschew this approach for two reasons.

First, measurement error in W_t introduces significant spurious correlation between the dependent and independent variables. A more sophisticated estimation technique is therefore required. One could employ a two stage procedure, instrumenting for A_t/W_t with A_t/Y_t (where Y_t is lifetime resources). In the results reported here, we simply substitute A_t/Y_t for A_t/W_t in the basic specification. Estimates based on instrumenting for A_t/W_t (not reported) differed very little from these results.

Second, data on g is extremely poor. Inference of g from successive observation of benefits received by the same individual is subject to enormous measurement error (due to variance in reporting). Alternatively, one might attempt to form an estimate of g based on the proportion of benefits which are unindexed. Presumably, this is closely related to the proprotion of benefits received from private sources (PROP), since government pensions (including Social Security) are indexed, while most private pensions are not. However, the accuracy of this estimate would be questionable, particularly since many apparently unindexed private pensions are de facto indexed by "good will" increases in benefits. Although one would nevertheless expect PROP and g to be negatively correlated (due to the lack of ubiquitous indexing), the magnitude of this correlation is unknown. The use of PROP to proxy for g would only allow us to judge the directions of various effects, rather than their magnitudes.

These considerations lead us to estimate the following modified version of equation (18):[47/]

$$(19) \quad \ln W_{t+1,i}/W_{t,i} = \beta_0 + \beta_1 \frac{A_{t,i}}{Y_i} + \beta_2 \, PROP_{t,i} \cdot \frac{A_{t,i}}{Y_i} + \varepsilon_{t,i}$$

where i indexes household. Rather than attempt to recover γ and r and to test parameter restrictions, we simply inspect the pattern of coefficients. For a sample dominated by dissavers (savers), β_1 should be negative (positive). Since PROP is negatively correlated with g, β_2 should be positive. We will, in addition, estimate a version of (19) where $\ln A_{t,i}/Y_i$ is substituted for $A_{t,i}/Y_i$. Since several levels of approximation have been used in deriving equation (19), we have no great attachment to any particular functional relationship; it is therefore important to determine whether or not the signs of estimated coefficients are sensitive to such alternative specifications.

Unfortunately, estimation of equation (19) may be contaminated by spurious correlation between PROP and ε. Individuals with large private pensions may, for example, be atypical (wealthier, less impatient). Alternatively, large values of PROP may reflect greater exposure to inflation risk, which would in turn have behavioral implications. We remedy these problems by including PROP as a separate right hand side variable in the estimating equation:

$$\ln W_{t+1,i}/W_{t,i} = \beta_0 + \beta_1 \frac{A_{t,i}}{Y_i} + \beta_2 \, PROP_{t,i} \frac{A_{t,i}}{Y_i} + \beta_3 PROP_{t,i} + \varepsilon_{t,i} \quad .$$

Our expectation is that the spurious effects described above will be captured in the estimated value of β_3: although there are many reasons to believe that PROP is systematically related to ε, it is much more difficult to explain why the partial correlation (controlling for PROP) between the interaction and error terms would be nonzero.

We estimated these specifications separately for single individuals and couples, using $t = 1975$ and $t + 1 = 1979$. The second period was chosen so that the samples would be more representative of typical retirees. Results are presented in Table 8 and 9.

Consider first the regressions for single individuals (Table 8). Specification 1 corresponds to equation (19). Referring to equation (18), we see that the estimated intercept measures the four-year (non-annuitized) dissaving rate. The particular value presented in Table 1 implies a yearly dissaving rate of about 6%, which is on the high end of the estimates presented in Section 3. Since those estimates were not corrected for annuities, this leads one to suspect that annuitization increased the rate of accumulation for this group, contrary to our theoretical predictions. The point estimate of the coefficient on A/Y confirms this suspicion; however, it is estimated very imprecisely, and a range of magnitudes entirely consistent with the theory are well within a single standard deviation. Finally, we see that the coefficient of PROP.A/Y is negative, and statistically significant at a high level of confidence. This is, of course, inconsistent with the theoretical implications outlined above.

Adding PROP to this regression (specification 2) changes none of the qualitative conclusions, and in fact increases both the magnitude and statistical significance of the estimated coefficient on PROP.A/Y. Evidently, spurious correlation between PROP and ε have the effect of biasing this coefficient upwards. Notice also that the coefficient of PROP is statistically significant--its inclusion in the regression is warranted.

The pattern of estimates using log A/Y is only slightly different. Although this alternative specification obscures the interpretation of the intercept, the signs of the remaining coefficients may again be revealing. As before, the separate effect of annuitization is estimated very imprecisely. Furthermore, when PROP is omitted (specification 3), the estimated coefficient of PROP.A/Y is positive, though statistically insignificant. However, the inclusion of PROP drives this coefficient significantly negative as before; furthermore, the inclusion of PROP seems warranted on statistical grounds (its t-statistic is approximately 4).

We turn now to the regressions for couples (Table 9). The intercepts in specification 1 and 2 suggest a small positive saving rate, roughly consistent with that estimated in Section 3. While one cannot reject the hypothesis that this term is negative, values lying within two standard deviations are consistent with, at most, a dissaving rate of 2% per year. We remarked earlier that couples may nevertheless have intended to dissave--the observed accumulation may have been due entirely to unanticipated housing price inflation during this period. If this is so, annuitization should depress the rate of accumulation for this group. The coefficients of A/Y reveal that exactly the opposite is the case. While these coefficients are not statistically significant at conventional levels, notice that these levels are surpassed by the estimated coefficients of log A/Y in specifications 3 and 4. Together, these estimates strongly suggest that annuitization increased accumulation rates for this group.[48] If so, there are two possibilities: either couples are intentional net savers after retirment (which requires us to accept somewhat implausible behavioral parameters to rescue the life cycle model), or the response among couples of saving to annuitization is inconsistent with life cycle motives.

Further evidence against the life cycle hypothesis is again generated by the estimated coefficients of PROP.A/Y and PROP.log A/Y. The pattern here closely resembles that for single individuals. In three of four specifications, the estimated parameter is negative; in two of these it is statistically significant at conventional levels. Once again, only specification 3 yields a point estimate consistent with theory. However, specification 4 reveals that the omission of PROP is unwarranted on statistical grounds.

Although we have reported relatively few regressions in this section, our estimates were quite robust with respect to the inclusion of other potentially important variables. Adding age of respondent, health, and number of living children did not, for example, substantively alter any of the results discussed above.

6. Conclusions

If, as suggested here, the pure life cycle hypothesis fails to account for savings behavior after retirement, then it is important to determine whether this behavior is consistent with other theories. One possibility is to maintain life cycle motives, while posing the problem of wealth accumulation within a different institutional setting. In particular, the models of Kotlikoff and Spivak [1981], and Bernheim, Shleifer and Summers [1984] portray intergenerational transfers as a mechanism for facilitating intrafamily exchange. Alternatively, one can supplement the life cycle model with a traditional bequest motive. Fortunately, these alternatives generate testable empirical implications. Bernheim, Summers and Shleifer present econometric and other evidence to support a strategic bequest motive. My own work in progress (preliminary results are presented in Bernheim [1984]) considers whether or not the data are also consistent with a model of household preferences augmented with intergenerational altruism.

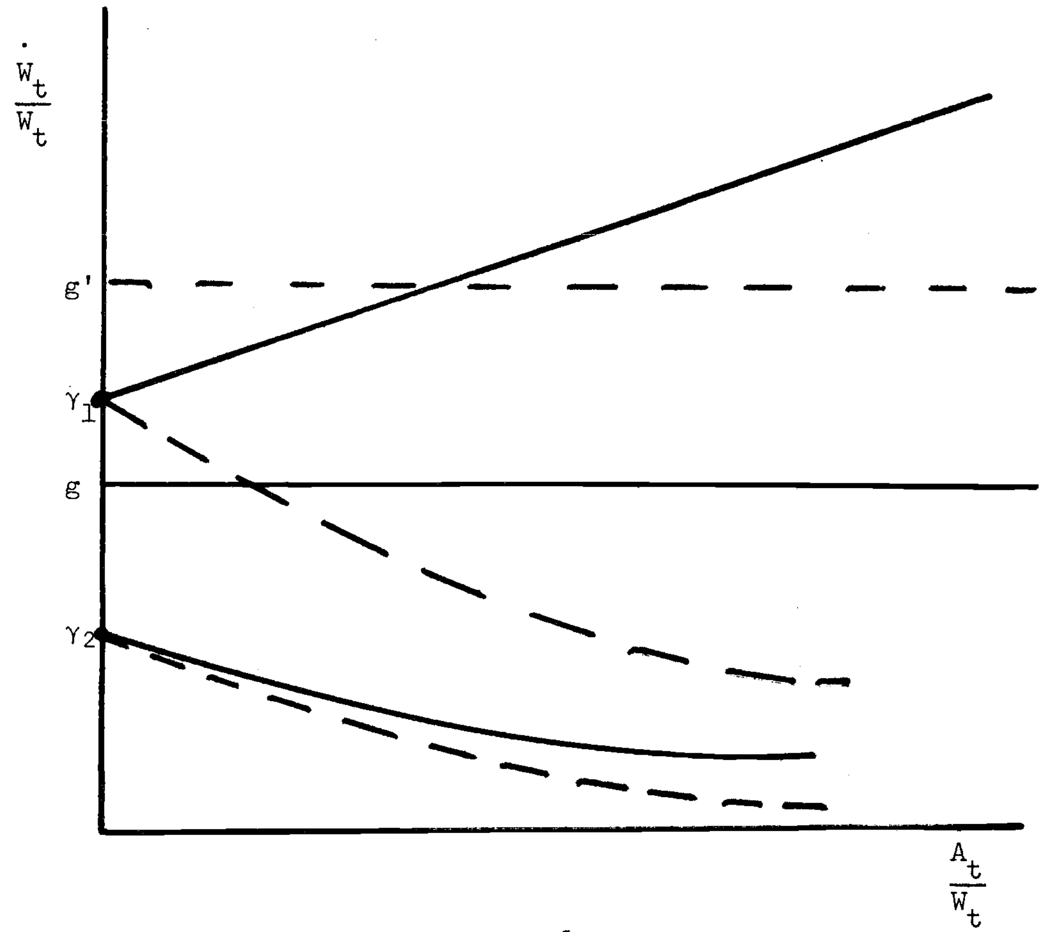

Figure 1

Footnotes

1/ These tests should not be confused with those of Feldstein [1974, 1977], Feldstein and Pellechio [1979], Kotlikoff [1979], and others who examine the effect of involuntary annuitization on levels of bequeathable wealth holdings.

2/ Notice that this definition does not include the value of durable goods. It is quite likely that, as a result, the data understate the true rate of dissaving (elderly individuals probably engage in few purchases of new durable goods, while old goods depreciate). The resulting bias is, however, likely to be small.

3/ Ferber et al. [1969] documents a tendency for misreporting of assets to be related to the respondent's level of wealth.

4/ This can often be inferred from the corresponding completion codes, or from the implausibility of recorded values.

5/ Presumably, if an individual does not own a home, it is straightforward to report 0.

6/ Earned income does not exceed $1000 per year in any year after retirement.

7/ Note that this group is not contaminated by any households which retired in the interim.

8/ Aware of this difficulty, Mirer re-estimates his regressions for the subsample of individuals who are over 75 years old. Although this does not completely eliminate the bias (in particular, many members of this subgroup may perform significant part-time work), and although this subsample may be dominated by outliers in the age spectrum, the robustness of Mirer's original estimates is suggestive.

9/ Lydall and Projector and Weiss simply group together all individuals over 65. Mirer reports that 37% of his sample is between 65 and 67 years old.

10/ Estimation using the 1969 wave yielded very imprecise estimates. However, it should be noted that the coefficient on age was slightly positive in all cases.

11/ Note that the samples sizes here are larger than those reported in Section 2. Since we employ cross-sectional data here, we do not insist that the households survive to a later sample year.

12/ Suppose $A_1 > A_2$. Then, if R is age of retirement,

$$\begin{aligned}
E(R|R \leqq A_1) &= \text{prob}(R \leqq A_2|R \leqq A_1)\, E(R|R \leqq A_2) \\
&\quad + \text{prob}(A_2 < R \leqq A_1|R \leqq A_1)\, E(R|A_2 < R \leqq A_1) \\
&\quad > [\text{prob}(R \leqq A_2|R \leqq A_1) + \text{prob}(A_2 < R \leqq A_1|R \leqq A_1)] \\
&\qquad E(R|R \leqq A_2) \\
&\quad = E(R|R \leqq A_2).
\end{aligned}$$

13/ To put it another way, the probability of living to 70 conditional upon surviving to 69 is higher for the average 60 year old who actually survives to 69 than it is for the average 60 year old in general since the latter sample includes relatively unhealthy people with low conditional survival probabilities who are likely to die before they reach 69.

14/ The secular rise in life expectancies may partially or completely offset this effect.

15/ Mirer attempts to correct only for the "cohort effect," and finds, not surprisingly, more striking evidence of positive saving during retirement.

16/ Mirer's procedure, in particular, seems seriously flawed: he regresses the ratio of net value in owner occupied housing to total net worth on age and total net worth. Elsewhere, he concedes that there is likely to be substantial measurement error in net worth. This builds in a strong, spurious negative correlation between the dependent variable and observed total net worth (as reflected by its negative coefficient and enormous t-statistic). Presumably, all coefficients in this regression are then estimated inconsistently.

17/ Mirer [1979], p. 439.

18/ In the first sample year, NLS respondents are 45 to 59, as opposed to 58 to 63 for the LRHS.

19/ Unfortunately, Diamond and Hausman do not report the total number of individuals retiring during their sample period. Their regressions were, however, based on approximately 1200 observations. Assuming a uniform distribution of age, only 400 would have reached 65 by the end of the sample period. This may in part account for the large standard error of their estimate. In contrast, the youngest LRHS respondent was 68 in 1979.

20/ Diamond and Hausman's sample will overrepresent early retirees. This may explain much of their findings; see the comments at the end of this section.

21/ Since early retirees are typically poorer and less healthy, this somewhat offsets the other effects.

22/ It would in any case be quite difficult to produce a "typical" sample, since the LRHS oversamples certain groups to begin with.

23/ Note that for the "retired in 1969" and "not retired in 1975" samples, no value is reported for bequeathable wealth in 1979, since we do not require household survival past 1975.

24/ This follows from smoothing of consumption.

25/ The net worth of workers fell by 3.4% ($2299) between 1969 and 1975, and rose by 11.6% ($8771) between 1975 and 1979.

26/ In a world without annuities, wealth would evolve in this way as long as preferences were homothetic.

27/ The law of large numbers requires the existence of certain moments.

28/ The assumption of independence deserves some attention. One might object that an individual who underreports assets in one year is likely to do so in the next as well. This creates no problems, as long as the fraction underreported by individual i does not change systematically with his wealth.

29/ Observations with larger β_i's will (given the same level of initial wealth) be more likely to remain in the sample.

30/ For example, over one-third of LRHS households retired in 1969 were single individuals; in 1975, this figure fell to one-sixth.

31/ Thomas Gustafson has pointed out that the data presented here are too aggregated to test this hypothesis--we cannot tell if the <u>same</u> households which sell businesses and property also become new homeowners during this period. In fact, this pattern might seem somewhat unlikely, since households which do not own homes often have virtually no other assets. Alternatively, the rise in average housing wealth may be primarily attributable to the purchase of more expensive houses by those liquidating business and other property holdings (new homeowners may have virtually no equity). Another possibility is that individuals who move at retirement typically discover that their current house is worth more than expected; the decline in other assets should then be counted as dissaving. By disaggregating the data, it should be possible to distinguish between these possibilities. This is left for future work.

32/ If the rate of time preference exceeds the discount rate, households will consume all resources immediately. If the inequality goes the other way, the transversality condition is violated, and no optimum exists. For equality, the household is completely indifferent between all consumption programs that exhaust his resources.

33/ See Yaari [1965].

34/ Households may still hold some bequeathable wealth if annuitization occurs through the family, as suggested by Kotlikoff and Spivak [1981]. It is, however, unclear whether one can infer anything from rates of dissaving in the context of their model.

35/ Implicitly, we assume that single year conditional survival probabilities are constant over time. In such a world, the actuarial value of an annuity does not change with age. In what follows, it should be clear that our central results do not depend upon this assumption. In particular, the argument which establishes that simple discounting is approximately appropriate depends only upon there being a relatively long interval before the nonnegativity constraint on bequeathable wealth binds. To take an extreme alternative, suppose death will occur at date T, with certainty. If an annuity contract promises to pay benefits past this date, those benefits are irrelevant. The appropriate value of an annuity (assuming either that the individual can borrow on benefits paid prior to T or that terminal benefits are not too large) is then just the simple discounted value of benefits, up to age T. In this very special case, actuarial valuation is exactly appropriate, and our technique (which includes benefits promised after T) is clearly in error. However, we have added the qualification that there must be a relatively long interval before the constraint on bequeathable wealth binds. Here, it binds as T, so if T is large, our method is, again, approximately appropriate. In general, however, if there is some maximum age, one could always improve our measure by excluding benefits promised after the maximum age.

36/ The transversality condition guarantees this inequality.

37/ This is consistent with the calculations in the next section. Previous studies have obtained lower estimates of annuitization ($A_0/(r - g)W_0$) specifically because they have employed actuarial valuation.

38/ For elderly individuals, single year survival probabilities are approximately 95%, so one can think of $\lambda = 0.05$ as representing the case where all discounting is due to uncertain length of life.

39/ While these calculations appear to confirm the superiority of simple discounting as a measure of value, the reader should bear in mind that any sample of elderly individuals may exhibit great behavioral heterogeneity. Thus, even if simple discounting is appropriate for the median household, it may be highly inaccurate when applied to rapid dissavers, who will reach a binding constraint quickly.

40/ Details are available from the author.

41/ If $\gamma > g$, the growth of bequeathable assets accelerates with annuitization. For this case, the non-negativity constraints never bind, and (17) is exactly appropriate.

42/ This requires an exceptionaly large amount of tedious algebraic manipulation. Details are avaliable from the author.

43/ Again, details are available from the author.

44/ We assumed inflation rates of 6% prior to 1969, rising to 9% by 1975, and 12% by 1979, remaining constant thereafter.

45/ While this conclusion appears warranted for the median household, we have ignored sample heterogeneity. This is particularly important, since rapid dissavers will reach a binding constraint on bequeathable wealth quickly, thereby rendering the use of simple discounting perhaps very inaccurate. Unfortunately, we cannot distinguish behavioral heterogeneity from measurement error.

46/ See Bernheim [1984] for a discussion.

47/ Note that since $PROP_{t,i} = P_{t,i}/A_{t,i}$ (where $P_{t,i}$ is private pension benefits), $PROP_{t,i} \cdot A_{t,i}/Y_i = P_{t,i}/Y_i$ (i.e., the $A_{t,i}$ terms cancell).

48/ This finding is confirmed by Diamond and Hausman [1980].

References

Atkinson, A.B. (1971). "The Distribution of Wealth and the Individual Life Cycle." Oxford Economic Papers, vol. 23, pp. 239-54.

Atkinson, A.B. and Harrison, A.J. (1978). The Distribution of Wealth in Britain. Cambridge University Press.

Bernheim, B.D. (1984a). "Annuities, Bequests, and Private Wealth." Stanford University (mimeo).

Bernheim, B.D. (1984b). "Life Cycle Annuity Valuation," Stanford University (mimeo).

Bernheim, B.D., Shleifer, A., and Summers, L. (1984). "Bequests as a Means of Payment," NBER Working Paper No. 1303.

Brittain, J. (1978). Inheritance and the Inequality of Material Wealth. Brookings Institution.

Davies, J.B. (1980). "Uncertain Lifetime, Consumption, and Dissaving in Retirement." Journal of Political Economy, vol. 89, pp. 561-78.

Diamond, P.A. and Hausman, J.A. (1980). "Individual Savings Behavior." MIT (mimeo).

Feldstein, M.S. (1974). "Social Security, Induced Retirement, and Aggregate Capital Accumulation." Journal of Political Economy, vol. 82, pp. 905-926.

Feldstein, M.S. (1977). "Social Security and Private Savings: International Evidence in an Extended Life-Cycle Model," in The Economics of Public Services. International Economics Association Conference Volume, M. Feldstein and R. Inman (eds.).

Feldstein, M.S. and Pellechio, A. (1979). "Social Security and Household Wealth Accumulation: New Microeconomic Evidence." Review of Economics and Statistics, vol. 61.

Ferber, R. et. al. (1969). "Validation of a National Survey of Consumer Financial Characteristics: Savings Accounts." Review of Economics and Statistics, vol. 51, pp. 436-44.

Fox, A. (1976). "Alternative Measures of Earnings Replacement for Social Security," in Reaching Retirement Age, Research Report No. 47, Office of Research and Statistics, Social Security Administration, Government Printing Office.

King, M. A. and Dicks-Mireaux, L. (1982). "Asset Holding and the Life Cycle." Economic Journal, vol. 92, pp. 247-67.

King, M.A. and Dicks-Mireaux, L. (1983). "Pension Wealth and Household Savings: Tests of Robustness." Journal of Public Economics, forthcoming.

Kotlikoff, L. (1979). "Testing the Theory of Social Security and Life Cycle Accumulation." American Economic Review, vol. 69, pp. 396-410.

Kotlikoff, L. and Spivak, A. (1981). "The Family as an Incomplete Annuities Market." Journal of Political Economy, vol. 89, pp. 372-91.

Lydall, H. (1955). "The Life Cycle in Income, Saving, and Asset Ownership." Econometrica, vol. 46, pp. 985-1012.

Loury, G. (1981). "Intergenerational Transfers and the Distribution of Earnings." Econometrica, vol. 49, pp. 843-867.

Mirer, T.W. (1979), "The Wealth-Age Relationship Among the Aged." American Economic Review, vol. 69, pp. 435-43.

Mulanaphy, J. (1974). Statistical Report: 1972-73 Survey of Retired TIAA - CREF Annuitants. New York.

Projector, D. (1968). Survey of Changes in Family Finances. Federal Reserve Board of Governors.

Projector, D. and Weiss, G. (1966). Survey of Financial Characteristics of Consumers. Federal Reserve Board of Governors.

Shorrocks, A.F. (1975). "The Age-Wealth Relationship: A Cross-Section and Cohort Analysis." Review of Economics and Statistics, vol. 57, 155-63.

Smith, J.D. (1975). "White Wealth and Black People: The Distribution of Wealth in Washington, D.C. in 1967," in his The Personal Distribution of Income and Wealth, New York.

Stiglitz, J.E. (1978). "Equality, Taxation, and Inheritance," in Personal Income Distribution, W. Krelle and A.F. Shorrocks (eds.). North-Holland, Amsterdam.

Yaari, M.E. (1965). "Uncertain Lifetime, Life Insurance, and the Theory of the Consumer." Review of Economic Studies, vol. 32, pp. 137-50.

Table 1

Wealth Level Regressions for 1975 Cross-Section

Variable	Singles All	Singles Retired	Couples All	Couples Retired
Constant	-10934	168757	34527	170171
	(36359)	(83408)	(37321)	(118587)
Age	379	-2442	65.6	-1925
	(593)	(1354)	(608.9)	(1930)
Y	0.0234	0.00892	0.0133	0.0196
	(0.0054)	(0.0134)	(0.0035)	(0.0102)
Sample Size	1605	213	5960	964

Table 2

Bequeathable Wealth by Year and Retirement Status

Variable	Single Individuals				Couples			
	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979
Bequeathable Wealth in								
1969	15008	30003	21743	48768	34818	66719	61304	87374
1975	11832	27183	20601	37943	26895	64420	60144	75709
1979	---	---	19209	37071	---	---	62610	84480
Fraction of sample with positive bequeathable wealth in								
1969	0.685	0.858	0.821	0.896	0.915	0.938	0.957	0.949
1975	0.626	0.846	0.902	0.885	0.768	0.932	0.962	0.961
1979	---	---	0.879	0.885	---	---	0.954	0.951
1969 & 1975	0.415	0.675	0.671	0.719	0.621	0.818	0.855	0.827
1975 & 1979	---	---	0.688	0.750	---	---	0.861	0.844
Sample Size	270	240	173	96	504	1120	864	411

Table 3

Changes in Bequeathable Wealth by Year and Retirement Status

Variable	Single Individuals				Couples			
	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979
log W_{75}/W_{69}								
Mean*	-0.198 (0.108)	0.113 (0.131)	0.256 (0.143)	0.025 (0.180)	-0.125 (0.066)	0.171 (0.038)	0.123 (0.043)	0.077 (0.063)
Median	-0.186	0.152	0.131	0.009	-0.086	0.181	0.149	0.170
Fraction < 0	0.580	0.444	0.457	0.507	0.527	0.381	0.407	0.418
log W_{79}/W_{75}								
Mean*	---	---	-0.285 (0.120)	0.021 (0.164)	---	---	0.028 (0.044)	0.095 (0.055)
Median	---	---	-0.104	0.176	---	---	0.074	0.133
Fraction < 0	---	---	0.546	0.375	---	---	0.415	0.403

* Estimated standard errors in parentheses.

Table 4

Breakdown of Bequeathable Wealth for Retirees

Type of Wealth and Year	Single Individuals Retired by 1969	Single Individuals Retired by 1975	Couples Retired by 1969	Couples Retired by 1975
House*				
1969	6122 (0.307)	4175 (0.260)	13700 (0.688)	15013 (0.627)
1975	6424 (0.322)	9893 (0.468)	13944 (0.581)	25481 (0.791)
1979	---	8268 (0.416)	---	28934 (0.775)
Business and Property				
1969	1312	12042	6172	29625
1975	914	4575	3401	14013
1979	---	4143	---	14966
Financial Wealth				
1969	7718	5790	15654	17635
1975	4646	6509	10119	21509
1979	---	6949	---	19076
Non-Mortgage Debt				
1969	143	263	709	969
1975	153	374	567	861
1979	---	192	---	366

* Percent owning a home is given in parentheses.

Table 5

Wealth Trajectories for Hypothetical Parameter Values*

Calculated Parameter	Assumed Value of λ	
	0.05	0.07
γ	-0.020	-0.040
T	42	27
ϕ	0.114	0.254
ψ	0.027	0.090
γ^r	-0.016	-0.026
γ^w	-0.052	-0.093

* For these calculations, we assumed $r = 0.03$, $g = 0$, $\alpha = 0$ (i.e., the logarithmic case), $A_0/(r - g)W_0 = 2$, and $t = 6$.

Table 6

Changes in Annuity Wealth by Retirement Status

Variable	Single Individuals				Couples			
	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979
Annuity Wealth								
1969	75002	80049	69885	64722	105699	95429	89340	94452
1975	73644	90580	77507	73048	100012	103500	95061	99762
1979	---	---	77131	72974	---	---	93910	100112
log A_{75}/A_{69}								
Mean*	-0.0095 (0.0042)	0.115 (0.0036)	0.093 (0.005)	0.102 (0.007)	-0.040 (0.005)	0.076 (0.003)	0.065 (0.003)	0.051 (0.006)
Median	0.006	0.147	0.114	0.147	0.002	0.080	0.067	0.068
Fraction < 0	0.311	0.032	0.019	0.023	0.442	0.090	0.141	0.164
log A_{79}/A_{75}								
Mean*	---	---	-0.004 (0.001)	-0.005 (0.005)	---	---	-0.011 (0.002)	-0.007 (0.003)
Median	---	---	-0.001	-0.001	---	---	-0.001	-0.001
Fraction < 0	---	---	0.963	0.943	---	---	0.904	0.802

* Estimated standard errors are in parentheses.

Table 7

Changes in Total Wealth by Retirement Status

Variable	Single Individuals				Couples			
	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979	Retired by 1969	Not Retired by 1975	Retired by 1975	Not Retired by 1979
Total Wealth								
1969	90009	110051	91600	113491	140516	162148	150643	181826
1975	85475	117763	98108	110989	126906	167920	155205	175471
1979	---	---	96340	110045	---	---	156520	183141
log TW_{75}/TW_{69}								
Mean*	-0.067 (0.027)	0.0611 (0.028)	0.071 (0.052)	-0.021 (0.065)	-0.094 (0.014)	0.062 (0.012)	0.055 (0.012)	0.029 (0.019)
Median	0.005	0.134	0.087	0.076	-0.044	0.097	0.071	0.065
Fraction < 0	0.450	0.228	0.237	0.337	0.624	0.302	0.328	0.370
log TW_{79}/TW_{75}								
Mean*	---	---	-0.046 (0.025)	0.011 (0.049)	---	---	0.013 (0.011)	0.023 (0.020)
Median	---	---	-0.003	0.004	---	---	0.005	0.027
Fraction < 0	---	---	0.586	0.467	---	---	0.479	0.424

* Estimated standard errors are in parentheses.

Table 8

Single Individuals, Retired by 1975
Dependent Variable: log W_{79}/W_{75}

Variable	Specification			
	1	2	3	4
constant	-0.235 (0.133)	-0.274 (0.131)	-0.609 (0.746)	-0.192 (0.706)
A/Y	0.031 (2.01)	0.227 (1.96)	---	---
Log A/Y	---	---	-0.076 (0.172)	0.017 (0.163)
PROP	---	8.47 (3.22)	---	-52.2 (13.3)
PROP•A/Y	-315 (95.2)	-735 (184)	---	---
PROP•log A/Y	---	---	0.425 (0.397)	-11.4 (3.04)

Table 9

Couples, Retired by 1975
Dependent Variable: log W_{79}/W_{75}

Variable	Specification			
	1	2	3	4
constant	0.0144 (0.0529)	0.0360 (0.0559)	0.531 (0.262)	0.763 (0.273)
A/Y	1.46 (1.20)	1.27 (1.21)	---	---
Log A/Y	---	---	0.113 (0.061)	0.165 (0.064)
PROP	---	-0.403 (0.339)	---	-3.55 (1.26)
PROP•A/Y	-25.7 (10.4)	-13.7 (14.3)	---	---
PROP•log A/Y	---	---	0.105 (0.054)	-0.665 (0.279)