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# The distribution impact of the social security program, 1962-1972

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THE DISTRIBUTION IMPACT OF THE SOCIAL SECURITY PROGRAM, 1962-1972

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# I. STATEMENT OF INQUIRY AND RESULTS

For nearly five decades, the social security program[1] has grown in scope, worker coverage, budgetary significance, and, until quite recently, popularity. However, the federal Old-Age, Survivors, Disability, and Health Insurance (OASDHI) program has entered a new phase in its long, convoluted history--a phase marked by public confusion, critical debate, budgetary insolvency, and controversy. This dissertation investigates a cause of the controversy, the income redistribution objective of the program. The old-age insurance portion of the social security program has two primary objectives: 1) to insure retirees against economic risk over an uncertain retirement period when potential earnings are low or zero; and 2) to redistribute income within an age cohort and across generations. The former objective alters the pattern of income receipts across the individual's life cycle, whereas the latter alters the distribution of lifetime income within an age cohort and across generations. Over time, policymakers have shifted the emphasis of the program away from traditional insurance principles, or "individual equity," toward a distribution of benefits based on the presumptive needs of retired persons and their dependents, or "social adequacy."

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[1]"Social security" is broadly defined as the federal Old-Age, Survivors, Disability, and Health Insurance (OASDHI) program. Prior to 1966 when the health insurance program was added, it was referred to as OASDI. This paper confines its analysis to the old-age (OAI) portion of OASDHI, which includes primary worker, spousal, transitional, and special age-72 benefits.

The primary, although not exclusive, emphasis of the program has become an attempt to extend a minimum standard of income security to all "effectively" retired persons in pursuance of social justice. The apparent dual nature of the program was not problematic until recently because taxes were kept at acceptable levels, covered retirees were generally net gainers, and, to a lesser extent, the program was conveniently cast in a traditional insurance-like framework. The first generation of OAI beneficiaries received exorbitant rates of return on prior OAI contributions owing to the fact that they had few years of coverage in the program and a relatively long benefit collection period. Subsequent generations have benefitted from the relative immaturity of the program, which made possible extremely low tax rates and frequent increases in benefit levels. As the system matures, meaning the contribution period eclipses the entire work history, the size of the intergenerational transfer will diminish. In addition, the probability of being a net loser will increase, drawing further attention to the cause of the potential loser-gainer scenario--the redistribution objective.[1]

This dissertation does not address the legitimacy of the redistribution objective; instead, it seeks to examine the program's effectiveness in redistributing income within and across retirement cohorts. Four interrelated issues are investigated: 1) Does the OAI portion of the

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[1]Parsons and Munro (1977) find that within the next 50 years the intergenerational transfer will disappear completely; hence, each retirement cohort will distribute amongst its members the amount of money they initially paid into the program. Similar results regarding the diminution of the intergenerational transfer were found by Freiden et al. (1976) and Burkhauser and Warlick (1981) (see Chapter III).

social security program redistribute income in favor of low-income beneficiaries? 2) Does the current OAI program redistribute benefits in favor of women, as a group, at the expense of their male counterparts? 3) How does the wife's work status affect the distribution of OAI benefits within and across family types? 4) Are spousal benefits distributed principally to needy, dependent spouses? Answers to the aforementioned questions are needed to assess the effectiveness of the current OAI program in satisfying its intented objectives and to shed light on inequities and inadequacies resulting from specific provisions in the law.

The distributional impact of the OAI program is isolated by "disentangling" or "decoupling" the insurance portion of the OAI benefits from the redistribution portion. The insurance disentanglement employs the actuarial standard of Burkhauser and Warlick (1981), whereby a retired worker's 1972 OAI benefit level is compared to the benefit level the worker would have received from purchasing an actuarially fair life annuity with his or her accumulated OAI contribution on the date of retirement. (Burkhauser and Warlick define this difference as the "transfer component.") The life-cycle framework devised by Burkhauser and Warlick is extended in this dissertation to account for the monthly disbursement of OAI benefits and price indexing. This approach allows us to measure the distributional effects of the progressive benefit formula, spousal benefits, and price indexing.

Chapter II presents a brief historical overview of the OAI program with emphasis placed on features of the law to be examined in this study.

Evidence from previous empirical studies investigating the distributional impact of the social security program are discussed in Chapter III. The life-cycle model and conditions for an actuarially fair retirement system are presented in Chapter IV. In Chapter V, the assumptions underpinning the model, the data set and sorting technique, computational formulas, annuity-type counterfactuals, and redistribution estimates are briefly explained. The generalised quadratic regression models by marital status, and a detailed discussion of the model variables are presented in Chapter VI. In Chapter VII, descriptive evidence and evidence from the estimation of the regression models are presented and interpreted. A summary and conclusions appear in Chapter VIII.

## II. HISTORICAL OVERVIEW

### A. 1935 Old-Age Insurance Program

The social security program in the United States is a dynamic federal income maintenance program, which has evolved over its brief 49-year history from a strictly worker-only retirement program to a full-fledged, comprehensive old-age social insurance program. The 1935 old-age program provided retirement benefits to covered workers only. Benefit levels were a function of total covered wages earned by the worker over her work history, and financed by a flat-rate payroll tax levied on the employee and employer. Although the OAI program was partially funded, it was not distributionally neutral. Initial benefit levels were determined by a mildly progressive benefit formula, and benefit payments were not adjusted to reflect different life expectancies of male and female beneficiaries. Hence, even in the early years of the program (prior to 1940, when the first benefits were paid) some redistribution within a cohort, though not across cohorts, was mandated.

### B. Spousal Benefit Provision

A major drawback to the initial program was its relative ineffectiveness in providing adequate income protection for dependents of covered workers, and soon-to-be and already retired workers. Incremental changes in benefit coverage and funding principles were introduced in the form of amendments to the Social Security Act of 1935 to enhance the effectiveness of the program in pursuing the goal of income adequacy for aged persons--the nation's most identifiable impoverished group. The 1939 amendments provided spousal and survivor benefits for women married to covered workers.

The 1937-1939 Advisory Council's recommendation for noncontributory, supplemental security benefits to wives and widows of covered workers was a conscious attempt to ameliorate the economic hardships imposed on this group of women because of the incidental retirement or death of the primary earner who, at that time, did not have sufficient earnings history to satisfy his own economic needs in retirement let alone those of his dependents.[1] The receipt and absolute size of the supplemental benefits were linked to the husband's earnings history, preserving the illusion of an insurance program. The supplemental benefits provided family protection, although contributions were based on an individual worker's employment and earnings history exclusively. The OAI program legislated in 1939, and to a large extent operating today, effectively subsidized the traditional family structure characteristic of that time period. It is, however, important to note that the Council's recommendation was reflective of the socio-cultural-economic milieu of that period.

The typical American family in the late 1930s was characterized by life-long marriages where the female assumed the primary responsibility

[1]Noncontributory, supplemental security benefits were not extended to husbands of female workers until 1950. In 1950, husband and widower benefits were extended to the husband of a female worker if he could prove that one-half of his support came from his working wife or deceased wife. The "dependency test" was stricken from the law after it was declared unconstitutional by the Supreme Court in 1977 (Califano v. Goldfarb).

for nuturance and home management and the male assumed the "breadwinner" role. Married women, as a group, had weak labor force attachments and, as a consequence, were disproportionately represented outside the labor force. (In 1939, only one out of four married women worked outside the home, and three out of 20 households had both husband and wife employed outside the home simultaneously.) Most women, therefore, lacked independent OAI protection. The presumption of dependency, on behalf of all women, was consistent with demographic characteristics and did eliminate a severe inadequacy present in the original version of the strictly worker-only retirement program.

The Council realized that in the near future, and especially in the distant future, married women would be dually entitled to both primary and spousal retirement benefits. The provision of overlapping benefits to married women as independent earners and dependent spouses was inconsistent with the intent of the noncontributory, supplemental security benefit provisions--protecting a needy group from economic hardship resulting from the "breadwinner's" retirement or death. To avoid the overlapping benefit problem, the dual-entitlement provision was introduced in conjunction with supplemental benefits as a variant of a means test. According to the dual-entitlement provision, if a married woman is entitled to two benefits simultaneously--primary and spousal (survivor)--she will receive the larger of the two benefits. The base of her benefits is her own primary benefit amount which is then augmented by the difference between her supplemental benefit and primary benefit amounts. The dual-entitlement provision was a noncontroversial addition to the program because it pertained to a small fraction of the entire beneficiary population, and it was consistent with the generally accepted social adequacy goal of the program.

As mentioned above, the provision of spousal and survivor benefits to women married to covered workers in accordance with the dual-entitlement rule was noncontroversial in light of the demographic characteristics of the 1930s, 1940s, and 1950s. However, as women, especially married women, increased their participation in the labor force, a greater proportion of female beneficiaries qualified for independent as well as dependent's benefits.[1] Since the dual-entitlement provision guarantees the dually entitled woman the larger of the two benefits, she must forego the other benefit to which she is entitled. The design of the program gives preferential treatment to dependent, nonworking married women vis-à-vis independent, working married women. A nonworking married woman receives dependent spousal benefits (equal to 50 percent of her husband's primary insurance amount (PIA)) at a zero marginal cost, whereas a working married woman receives dependent spousal benefits at a marginal cost equal to her total OAI contributions, or primary worker benefits at a marginal cost equal to 50 percent of her husband's PIA. The working married woman may, either totally or fractionally, duplicate protection already afforded to her when classified as a dependent on her

[1]There was a 20-fold increase from 1950 to 1971 in women receiving primary-worker benefits. At the end of 1971, there were 23 million adult beneficiaries of which 13.8 million were women. Fifty percent of the female beneficiaries were receiving primary-worker benefits and 50 percent were claiming auxiliary benefits. The average monthly check for female beneficiaries was $100 (Bixby, 1972).

husband's account. Hence, the dual-entitlement provision acts as an implicit tax on the working married woman, since she receives only marginal accretions to her benefit level in return for her contributions into the program. The dual-entitlement provision implicitly penalizes the working woman for seeking financial independence and subsidizes the financial dependency of the nonworking married women. The effect of the dual-entitlement provision may, especially in light of legislated increases in the payroll tax and the relatively low earnings potential of most females, have an increasingly severe work-disincentive effect, and, in addition, may erode the progress women, as a group, have made in achieving financial liberation.

In addition to generating inequities across married women who have made different labor-homemaker decisions, the provision of noncontributory, supplemental benefits generates inequities across household types, depending on marital status and the division of earnings within the household. A two-earner household with equal earnings (a household where the husband and wife are gainfully employed outside the home) will receive lower combined benefits relative to a one-earner household (a household where either the husband or wife is gainfully employed outside the home) if the combined earnings of the two-earner unit is less than the taxable maximum for a single-earner. A two-earner household receives higher benefits compared to a one-earner household when their combined earnings are greater than the taxable maximum for a single earner; however, the two-earner couple pays more in the form of contributions to receive the higher benefit level (Bixby, 1972). The inequities between the one-earner and two-earner households have become more pronounced in light of the historic four-decade upswing in the employment participation of women.

Single persons, of either sex, are placed in a strategically inferior position in a retirement program that provides family protection based on an individual worker financing scheme. Single households are assigned the same tax liability as married households; however, the married household is afforded a greater package of benefits. Single and married workers are treated equally on the contribution side of the program, but they are treated as unequals on the benefit side since the married household is eligible for dependent benefits not similarly extended to a single person.

The inequities resulting from the 1939 amendments may, at first blush, appear justified in light of the social adequacy objective. However, the features of the program and the incidental inequities must be juxtaposed to modern demographic characteristics to ascertain whether or not the actual effect of the law is consistent with its intent. The payment of spousal benefits presumes the financial dependency of the married woman and a traditional family structure. The traditional 1939 family does not typify the family of the 1980s or of the future. The modern family is characterized by interdependency rather than dependency. That is, the typical family today is an interdependent economic unit in which partners, of either sex, have occupational choice and, to a large extent, are not forced to assume sterotypical roles mandated by societal norms. Women, as a group, are exercising their right to occupational choice and seeking covered employment outside the home.[1] This protracted trend will intensify the inequities among women who have made different labor-homemaker decisions. These inequities are a direct result of noncontributory, supplemental security benefits coupled with the dual-entitlement provision.

There remains a shrinking proportion of women who choose to be homemakers and, therefore, may need income protection in their retirement years.[2] According to the OAI program, the group of modern-day homemakers is presumed to be an identifiably needy group. Information on the pattern of lifetime work for married women is incomplete; however, most empirical evidence suggests that there is an inverse relationship between family income (net of the wife's earnings) and a wife's labor force participation (Boskin, 1973; Cain, 1966; Garfinkel and Masters, 1977). This evidence suggests that the homemaker choice is a more viable option for high-income families, which would tend to refute the needy-group argument supporting the provisions of noncontributory, supplementary benefits. Holden (1979), using a single-period analysis, found that supplemental benefits were disbursed proportionately to couples in all income categories. Thus, spousal benefits were being distributed to

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[1]In 1940, 17 percent of married women were represented in the labor force compared to 52 percent in 1981. The labor force participation of women is expected to continue its upward trend in the future. The actuaries of the Social Security Administration project a labor force participation rate of approximately 67 percent in 1990 for women age 25 to 54.

[2]In 1984, only six percent of all families were made up of the traditional nuclear family where the man works and the woman is a full-time homemaker.

spouses who were not needy according to poverty standards. This issue is addressed in a life-cycle context to determine if supplemental benefits adequately serve the 1939 objective of protecting a group of aged persons experiencing economic hardship. In addition, sex differentials in survivorship are employed to determine if women, as a group, are made differentially better off relative to their male counterparts since OAI benefits are not adjusted to account for different life expectancies between men and women of the same age.

### C. Progressive Benefit Formula

Traditional insurance funding principles were abandoned in 1939 for deficit financing, or what is more commonly referred to as "pay-as-you-go" financing. The deficit financing provision mandated intergenerational transfers from the currently working population to the retired, nonworking population.[1] The disbursement of benefits to retired persons was based on a progressive benefit formula. The formula has become slightly more progressive over time.

The OAI program, by design, favors low-income households through the retirement benefit formula used to determine the worker's primary

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[1] The "Ponzi-like" financing scheme is financially sound provided economic and population growth exceed the growth in the size of the retirement population (Pechman et al., 1968; Samuelson, 1958).

insurance amount (PIA) from her average monthly earnings (AME).[1] The retirement benefit formula is structured to pay higher marginal and average benefit rates as the benefit base (AME) decreases. Therefore, the replacement rate (the ratio of retirement benefits to preretirement earnings) is higher for low-income households relative to high-income households. But high-income households receive more cash benefits per month in absolute dollars. The original OAI benefit formula was mildly progressive. The formula applied to average monthly earnings limited to $250 and paid 40 percent of the first $50 plus ten percent of the next $200. This formula has been periodically revised to favor low-income households. In 1972, the formula paid 108.01 percent of the first $110 plus 39.39 percent of the next $290 plus 36.71 percent of the next $150 plus 43.15 percent of the next $100 plus 24 percent of the next $100 plus 20 percent of the next $250.

This study examines the distributional impact of the progressive benefit formula to ascertain whether, in fact, low-income beneficiaries receive preferential treatment in the disbursement of benefits *vis-á-vis* high-income beneficiaries. The progressivity of the benefit formula has

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[1]The average monthly earnings is a summary measure of the worker's earnings history calculated by summing the total taxable earnings in the computation years divided by the number of months in the computation period. The AME measure was replaced by a wage indexed base called the average indexed monthly earnings (AIME) in 1977. The AIME indexes the worker's earnings so that taxable earnings earned at different points in the life-cycle are expressed in terms of the overall earnings levels prevailing in the year of eligibility. The PIA is the basis for all benefit payments and is a function of the worker's AME (or AIME after 1977).

been disputed because of empirical evidence suggesting that socio-economic characteristics influence life contingencies.

### D. Actuarial Reduction for Early Retirement

The actuarial reduction in the monthly benefit amount payable on entitlement applies to retired workers and dependents aged 62 to 64. The intent of this provision was to equalize the total actuarial value of benefits received by the beneficiary independent of the age of retirement. In 1956, provisions were added to the law permitting female beneficiaries to accept retirement benefits at age 62. If the female beneficiary applies for early primary benefits (in advance of age 65), her PIA is reduced by 5/9 of one percent per month under age 65 (maximum reduction of 20 percent). Dependents' benefits are reduced by 25/36 of one percent per month under age 65 with a maximum reduction cap of 25 percent. Identical provisions were extended to male beneficiaries in 1961.

### E. Delayed Retirement Credit

The benefit level (PIA) is adjusted upward if the primary beneficiary elects to retire after age 65. Like the actuarial reduction provision, the accretion feature was intended to equalize the actuarial value of the benefit stream independent of the age of retirement. As of 1972, a covered worker's benefit level was adjusted upward if she remains actively employed and she does not accept retirement between ages 65 and 72. Benefits were increased by 1/12 of one percent for every month the covered worker postpones retirement after age 65.[1] Accretions in benefit levels are truncated at age 72. This adjustment in benefits for delaying retirement is less than the actuarial adjustment for the shorter life expectancies of older beneficiaries; hence, the postponement of retirement is translated into a real loss in benefits over the remaining life span.[2]

## F. Earnings Test

The earnings or retirement test is a type of means test which reduces benefits to beneficiaries who continue to work past the age of 65. An earnings test has been in effect since 1935. According to the 1935 earnings test, all retirement benefits would be withheld if the beneficiary received <u>any</u> labor earnings during retirement. The extortionate nature of this test was, however, relaxed prior to the payment of the first benefits in 1940. The 1939 version of the earnings test permitted labor-related earnings up to $15 per month without the loss of retirement benefits; however, all benefits were forfeited if earnings exceeded $15. Since 1939, the earnings limit has been augmented periodically.

In 1972, retirement benefits were reduced if the beneficiary remained employed after receiving retirement benefits and her earnings

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[1]In 1983, workers who postpone applying for retirement benefits receive benefits that are increased by three percent for each year acceptance is delayed past age 65 up to a maximum of 15 percent.

[2]The loss in benefits may be partially or fully offset by the worker's higher PIA as a result of the worker's extended earnings history.

were in excess of 19 percent of the annual taxable maximum. Benefits were reduced by one dollar for every two dollars of post-retirement earnings between $1,680 and $2,880, but benefits were reduced by one dollar for every dollar of earnings above $2,880.[1] However, benefits were not reduced for worker-beneficiaries who were 72 or older in 1972.[2]

From a policy point of view, the earnings test is consistent with the basic purpose of social security, which is to fractionally replace lost earnings because the aged worker retires from the labor force. But, from the beginning, the earnings test has been controversial and strongly criticized. The "$1 for $2 and $1 for $1" withholding rate (or "$1 for $2" withholding rate since 1973) has been criticized because the withholding rate applies to labor income only (excluding nonwork income sources like dividends, rents, and pension payments) and for discouraging healthy older persons from seeking gainful employment in the market. The burden of the 50-to-100 percent withholding rate falls heaviest on the low-income aged because of their greater reliance on social security and employment earnings for financial security during retirement. Studies of the financial holdings of the aged show that most low-income persons do not have access to private pensions, private insurance, savings, and

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[1]Since 1973, benefits were reduced by one dollar for every two dollars of earnings above the earnings ceiling. Beginning January 1, 1983, worker-beneficiaries age 65 to 70 lost one dollar of benefits for every two dollars of earnings over $6,600 ($550 per month), whereas younger retirees, age 62-64, forfeited one dollar of benefits for every two dollars of earnings over $4,290 ($410 per month). Both earnings limits are automatically indexed.

[2]Beginning in 1983, the earnings test applies only to worker-beneficiaries who are 65 to 70.

other nonwork income sources to augment their retirement benefits (Freidman and Sjogren, 1981; Murray, 1972; Sherman, 1973). Most evidence suggests that the financial status of low-income persons remains unchanged at the outset of retirement in spite of "social security" for several reasons: 1) retirement benefits only partially replace employment earnings; 2) retirement benefits are reduced if the retiree has supplemental post-retirement earnings in excess of the earnings ceiling; and 3) low-income persons generally have insufficient nonwork income sources.

### G. Cost-of-Living Adjustment

In the mid-1960s, influential persons in Congress and the executive branch began to push for a bigger role for social security as an income source for the elderly. Congress approved benefit increases of 15 percent in 1969, ten percent in 1971, and 20 percent in July of 1972. In October, 1972, Congress passed the Social Security Amendments of 1972. The major features of this legislation were provisions for indexing the wage base used in computing initial benefits and for using the consumer price index to adjust payments to current beneficiaries. Although automatic indexing was legislated in 1972, it did not become effective until 1975. Legislated increases were substituted for automatic indexing in 1973 and 1974.[1] Benefits paid to current beneficiaries are annually

[1]The expansion of social security beginning in 1969 is described in Martha Derthick, Policymaking for Social Security (1979).

indexed whenever the consumer price index (CPI) rises by more than three percent.[1]

The social security retirement system is intended to insure beneficiaries against the economic risk of longevity. Indexing of benefits enhances this form of insurance in an inflationary environment. Because women, as a group, have a longer life expectancy than men, they receive on average more benefits from indexing. Indexing of benefits for retired workers keeps intact the relative benefit structure, since all benefit streams are adjusted by the same index.

[1]Benefits are adjusted annually if the CPI changes by three percent or more. If the CPI changes by less than three percent in a year, benefits will not be indexed until the cumulative change exceeds three percent.

# III. LITERATURE REVIEW

Although the objective characteristics of the OAI program, including the extent of insurance protection, have changed over time, its initial intent of providing adequate protection against long-term uncertainties associated with the cessation of labor force participation because of old age has remained undiminished. Specific features that have been added to the program over time, compromising its 1935 insurance principles, ultimately influence the estimated size of the redistribution component. However, the gradual shifting towards social adequacy has engendered biases in the program's operation. The alleged biases include the preferential treatment of women, traditional family structures, low-income households, and nonworking persons age 65 to 71. A more subtle, but no less important, bias incidental to the program concerns differential survivorship. Mortality studies indicate that specific socioeconomic characteristics influence survivor probabilities (Antonovsky, 1972; Gove, 1973; Kitagawa and Hauser, 1973; Metropolitan Life, 1975).[1] In a retirement program that pays benefits for the duration of life, persons with lower survivor probabilities (or shorter life expectancies), as reflected by specific, identifiable socioeconomic

[1] Survivor probabilities measure the likelihood of an individual life age x surviving to life age x+1.

factors, subsidize persons with relatively higher survivor probabilities (or longer life expectancies).[1]

The effects of the aforementioned biases (program- and worker-specific) have been investigated in numerous empirical studies using different data bases (representative individual and individual case history approaches), model assumptions, equity measures, and program and fairness definitions. However, independent of the methodology employed, virtually all empirical studies indicate that social security beneficiaries retiring prior to 1975 received above-normal rates of return on their contribution dollars, independent of income classification and other socioeconomic characteristics (Aaron, 1974; Brittain, 1972a; Burkhauser and Warlick, 1981; Campbell and Campbell, 1967; Chen and Chu, 1974; Freiden, Leimer and Hoffman, 1976; Okonkwo, 1976; Ozawa, 1974). Although there is consensus on the "money's-worth" issue, there is less agreement concerning the overall progressivity of the program (Aaron, 1974; Freiden, Leimer and Hoffman, 1976; Okonkwo, 1976; Ozawa, 1974).

Analysts using single-period methodology have acknowledged the OAI program as being the most effective U.S. government program in redistributing income to an impoverished group (Danziger, 1977; Danziger and Plotnick, 1975; Lampman, 1971; Ozawa, 1974). The cross-sectional findings purporting the "success" of the program, in terms of decreasing

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[1]The effect of socioeconomic factors on mortality is more pronounced for persons aged 25-64; however, the effect of these characteristics is still relevant, in most cases, at advanced ages (i.e., age 65 and older).

income inequality across all income classes, can be explained by several factors. First, the progressive benefit formula replaces a larger percentage of the low-wage earner's preretirement earnings than for the high-wage earner. The redistributive function of the formula would tend to reduce post-retirement income differentials within a retirement cohort, ceteris paribus. Second, a large percentage of the aged is eligible for retirement benefits. The blanket coverage of the program enhances the income position of all income classes within a retirement cohort and improves their income standing relative to the working population. The third factor pertains to the absolute size of the transfers to the aged. Public assistance is considered to be the most economically efficient program of all income-maintenance programs; however, social security, while being economically less efficient, has the greatest redistributive impact. This apparent disparity between economic efficiency and redistributive impact is best explained by the following analogy: a 100 percent share of a peanut is still a peanut, but a 50 percent share of an elephant is half an elephant. That is, the amount of total benefits received by the targeted population depends on economic efficiency and the total amount of the outlay. In 1971, social security (OASDI) and railroad retirement programs paid out $39 billion in cash benefits compared to cash benefits totalling $10 billion under public assistance. The last factor to be discussed concerns the use of cross-sectional methodology to assess the performance of a life-cycle program. Cross-sectional investigations into the performance of the social security program assess the redistributive impact of the program by examining the degree of income inequality before and after the payment of retirement benefits. Clearly, this approach fails to account for the "income smoothing" function of the program; hence, it tends to overstate the redistributive impact of the program.[1] Results derived from the single-period analyses are strongly disputed by researchers using life-cycle models of the OAI program.

Many researchers have investigated the effect of the social security program (OAI, OASI, and OASDI) on the distribution of lifetime income within a life-cycle framework. The distributional impact has been measured in terms of lifetime internal rates of return, lifetime contribution-benefit ratios, and Burkhauser-Warlick-type "transfer" components (initial OASI benefit levels less the benefit received from a life annuity purchased with the worker's accumulated OASI contributions on the date of retirement). The absolute size of the distributional impact measure has been found to be sensitive to specific identifiable factors, such as date of retirement, marital status, sex, race, income class, education level, and age at entry and retirement. The empirical estimates of redistribution also depend on the richness of the data base and the model assumptions regarding benefit inclusion, payroll tax shifting, life expectancy tables, and market interest rates. Several of the major findings from studies using each measure are discussed below.

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[1]The "income smoothing" feature of the program focuses on the transfer of labor earnings from the worker's high earnings years to her retirement years through the contribution-benefit mechanism of the program.

Studies investigating the extent to which the social security program redistributes lifetime income among subgroups of an age cohort using an internal rate of return measure have generally found that the internal rate of return on OAI (OASDI) contributions is negatively related to income, date of retirement, age at retirement (relative to age 65), education level, and positively related to age at entry. Internal rates of return were also found to be higher for women, nonwhite races, and married persons. Furthermore, rates of return were found to be higher for all subgroups the less the assumed backward shifting of the employer's share of the payroll tax. Similarly, the absolute size of the rate of return for specific socioeconomic groups varied depending on the extent to which life expectancy tables were disaggregated. Also, real internal rates of return were found to be significantly smaller than nominal rates, where the gap between the real and nominal measures increased the larger the inflation rate relative to the annual rate of growth of retirement benefits.

The most comprehensive studies using the internal rate of return measure have been conducted by Okonkwo (1976) and Freiden et al. (1976). Okonkwo, using longitudinal age-earning profiles estimated from four successive U.S. population censuses and life expectancies disaggregated by sex, race, and education levels, found that the internal rates of return were higher for couples relative to single persons, higher for nonwhites relative to whites, and higher for households located in the south as opposed to the north. He also found that rates of return varied inversely with education level; specifically, workers with eight years of schooling received the highest return and workers with 16 or more years of schooling received the lowest internal rate of return, independent of race, marital status, region, sex, or type of tax (OAI, OASDI). However, the degree of redistribution, measured by the gap between the internal rates of return across education levels, for the white subgroup is reduced by the longer life expectancies for white persons with more education; hence, the degree of progressivity (attributable to the 1974 benefit formula) was weakened when adjusting for the larger survivor probabilities for whites with more education. In conclusion, Okonkwo argues that the social security program redistributes income to blacks and low-income whites as intended by the law, but that the redistribution effect is dampened by the differential survivorship probabilities. Aaron (1974), on the other hand, finds that differential mortality rates fully offset the progressivity built into the retirement benefit formula; hence, the redistribution flow is reversed, having a perverse effect on the distribution of lifetime income.

Freiden et al. (1976), using the Continuous Work History Survey and survivorship probabilities disaggregated by age, sex, and race, found the OAI program to be "very" progressive. That is, internal rates of return were found to be significantly higher for low-income subgroups relative to high-income subgroups. Like Okonkwo, Freiden et al. found that women received higher real rates of return than men, everything else equal.

Other studies have estimated contribution-benefit (C-B) ratios and "transfer" components to measure redistribution. These studies generally support the findings of the studies employing internal rates of return.

The C-B studies show that the C-B measure is negatively related to the market interest rate used to accumulate contributions and discount benefits (Brittain, 1972a; Chen and Chu, 1974). Burkhauser and Warlick (1981), using the 1973 Exact Match File to estimate annuity-type "transfer" components, found that all income classes in the 1972 retirement cohort received positive benefit transfers from the OAI program. In addition, they found that the amount of redistribution, measured in absolute dollars, was roughly equal for high- and low-income subgroups. The middle-income subgroup received the largest transfer from the program overall.

The higher rates of return associated with marital status, date of retirement, age at retirement, and income can be explained by the program's design in conjunction with differential survivorship probabilities. Other factors influencing the size of OAI returns, such as sex, race, and education, can be explained by differential survivorship probabilities.

The higher returns associated with marital status are attributable to two independent factors: 1) the OAI program, by design, subsidizes the traditional (one-earner) family structure through the provision of spousal benefits in accordance with the dual-entitlement rule; and 2) married persons, independent of race and sex, have longer life expectancies, on average, than their nonmarried, divorced, or widowed counterparts.

The first factor is related to the program's design whereby a nonworking married person receives dependent spousal benefits (equal to

50 percent of the spouse's primary insurance amount (PIA)[1]) at a zero cost, whereas a working married person receives dependent spousal benefits at a cost equal to her total OAI contributions or primary worker benefits at a cost equal to 50 percent of her spouse's PIA. Recall according to the dual-entitlement provision, a person entitled to two benefits simultaneously will receive the larger of the two benefits, but must forego the other benefits to which she is entitled. A similar partiality towards married couples is exposed when single persons are compared to married persons claiming dependents' benefits with the same prior contributions. The single person receives a lower rate of return on her (his) initial OAI contributions relative to a married person collecting dependents' benefits with the same OAI contributions, since a married person is eligible for dependents' benefits not similarly extended to a single person without dependents. Burkhauser (1979), using data from the 1973 Exact Match File, found that one-earner married couples fare better than either two-earner married couples or single individuals because one-earner households receive spousal benefits at no additional charge, and single persons are forced to participate in a retirement system designed for married persons.

The second factor pertains to the longer life expectancy of married persons. Mortality studies conclusively show that married persons of each race and sex have longer life expectancies than nonmarried, divorced, or widowed counterparts (Gove, 1973; Kitagawa and Hauser, 1973;

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[1]The primary insurance amount is the amount payable to a retired worker who begins to receive retirement benefits at age 65.

Metropolitan Life, 1975). It is interesting to note that the differences between married and unmarried statuses are much greater for men than for women. For instance, white, single males age 65 and over experienced mortality 44 percent greater than the level of white, married males comparably aged. Similarly, white, single females age 65 and over have mortality levels nine percent higher than comparably aged white, married females. A single person has a shorter life expectancy, on average, relative to a married person of roughly the same age, everything else equal. Both of these factors taken together exert upward pressure on the rate of return on OAI contributions for the traditional family structure relative to the nontraditional family structure, although married persons, one-earner or two-earner, fare better than single persons.

The date-of-retirement factor reflects the relative maturity or immaturity of the retirement program.[1] The first generation of OAI retirees received exorbitant rates of return on their prior OAI contributions owing to the fact that they had few years of coverage in the program and a relatively long period of benefit collection. Subsequent generations have benefitted from the relative immaturity of the program, which made possible extremely low tax rates and frequent increases in benefit levels. As the system matures, meaning the contribution period eclipses the entire work history, the size of the intergenerational transfer and subsequent rates of return on prior OASI contributions will diminish. Parsons and Munro (1977) contend that within the next 50 years

[1]It takes approximately 40 years for a retirement program to reach full maturity.

the intergenerational transfer will disappear completely; hence, each retirement cohort will distribute amongst its members the amount of money they initially paid into the program. Fraiden et al. (1976) studied the retirement cohorts from 1967 through 1970 focusing on worker-only beneficiaries. Although all coefficients were small, they found that the 1968 retirees' rates of return were 2.27 percent higher than the 1967 retirees', whereas the 1969 retirees' rates of return were 1.76 percent lower than the 1967 retirees'. There was no significant difference found between the rates of return for the 1970 and 1967 retirees. Burkhauser and Warlick (1981) found a general decline in the percentage of redistribution over time. By dividing the 1972 cohort into three age cohorts: 66-67, 72-75, 81-85, they found that the oldest age cohort received the largest intergenerational transfers and that the youngest age cohort received the smallest.

Income is an important factor in determining the overall progressivity of the OAI program. The program, by design, favors low-income households through the retirement benefit formula used to determine the worker's PIA from her (his) average monthly earnings (AME).[1] The retirement benefit formula is structured to pay higher marginal and average benefit rates as the benefit base (AME) decreases. Therefore, the replacement rate (the ratio of retirement benefits to preretirement earnings) is higher for low-income units relative to high-income units,

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[1]The average monthly earnings is a summary measure of the worker's earnings history calculated by summing the total taxable earnings in the computation years divided by the number of months in the computation period. The PIA is a function of the worker's AME.

although high-income units receive more cash benefits per month in absolute dollars. Most studies have found the OAI program to be progressive. Fraiden et al. (1976) estimated the elasticity of the internal rate of return with respect to lifetime income for OAI benefits of -.278. Other studies using a broader definition of benefits and more disaggregated mortality rates have shown less progressivity than the Freiden et al. study (Okonkwo, 1976; Aaron, 1974).

The age-at-retirement factor influences the size of the return because of the early-retirement and delayed-retirement features of the program. Persons who choose to remain employed between the ages of 65-72 receive additional retirement benefits according to the number of incremental months employed during this age period. The PIA is increased by 1/12 of one percent for each month retirement is delayed after age 65 with a maximum adjustment of seven percent if the worker should remain employed until age 72. The accretion to the PIA, however, understates the shorter life expectancy of the worker who delays retirement. Alternatively, the actuarial reduction in the PIA for early retirement (retirement age of 62 to 64) is excessive. Freiden et al. (1976) found that the optimum age at retirement, in terms of maximizing the internal rate of return, is 65.

The last factor to influence the rate of return or extent of redistribution is differential mortality rates. The Kitagawa and Hauser study (1973) on differential mortality rates in the United States indicates that socioeconomic factors, especially sex, race, occupation, income, eduction, and marital status, influence the individual's probability of dying at (surviving to) a specific life age. The effect of marital status on survival was mentioned earlier and, hence, will not be discussed further. Mortality rates were found to be negatively related to income and education, which elicit the opposite effect of the progressive benefit formula on rates of return. Mortality rates were also found to be higher for men relative to women and nonwhites relative to whites. Hence, women, on average, can expect to receive a higher return on their OAI contributions vis-á-vis male counterparts given that, ceteris paribus, women have, on average, longer life expectancies than men. Freiden et al. (1976) found that women can expect rates of return on their OAI contributions that are approximately 8.8 percent higher than men, everything else equal, and that nonwhites can expect rates of return approximately 1.9 percent lower than whites.

# IV. THE LIFE-CYCLE MODEL

To evaluate the redistribution of the OAI program, the program was be divided into two flows of money--an inflow of contributions and an outflow of benefits. During the worker's earning years she pays in a flow of contributions, in the form of a flat-rate payroll tax, earmarked for the OAI program in exchange for a promise of a steady stream of real income in the latter phase of her life cycle. The accumulated value of the worker's contribution, $TC_i$, paid in over the work history is determined using a traditional compounding scheme and a nontraditional roll-over compounding scheme.

The traditional compounding scheme calculates the total OAI contributions, $TC_i^1$, credited to the covered worker's account on the date of retirement by

$$TC_i^1 = \sum_{y=B}^{RE} T_{yi} \prod_{j=y}^{RE} (1+r_j) \tag{4.1}$$

where $T_{yi}$ = OAI contributions in year y for individual i,

$r_j$ = annual yield on U.S. government bonds in year j,

RE = year of retirement, and

B = first year in covered employment.

The nontraditional roll-over scheme calculates the total OAI contributions, $TC_i^2$, by the generalized form of

$$C_{yi} = T_{yi}(1+r_w)^w (1+r_z)^z (1+r_x)^x \tag{4.2}$$

where $C_{yi}$ = compounded value of individual i's contributions paid in year y in the retirement year,

$r_w, r_z, r_x$ = appropriate U.S. bond rates, and

$w, z, x$ = bond maturities.

The value of total OAI contributions, $TC_i^2$, is calculated by adding together the compounded annual contributions, or

$$TC_i^2 = \sum_{y=B}^{RE} C_{yi}. \tag{4.3}$$

Annual contributions are carried through time according to a bond roll-over scheme. That is, it is assumed that the government invests the full amount of the worker's yearly OAI contributions, credited to her account as of the end of the year in question, into a government bond with the longest maturity that does not exceed the number of years from the date of investment to retirement. The coupon and principal are rolled over immediately upon maturity into the next longest bond that has a maturity period no longer than the difference between the roll-over date and the expected date of retirement. The superscripts w, z, and x reflect different bond maturities and sum to the number of years from year B to the retirement date.

The value of OAI contributions, $T_{yi}$, in equations (4.1) and (4.2) depends on the year the income is earned, y, the amount of income earned, $E_{yi}$, relative to the maximum taxable earnings base, $M_y$, and the relevant OAI tax rate, $t_y$. The individual's taxable earnings for the years 1937-1950 was determined by

a) $T_{yi} = t_y E_{yi}$ when $E_{yi} \leq M_y$ x; and

b) $T_{yi} = t_y E_{yi} + 1/2 t_y (E_{yi} - M_i)$ when $E_{yi} > M_y$.

For the years 1951-1954, taxable earnings were determined in three different ways depending on the type of income earned and the relationship between income earned and the maximum taxable earnings base. In the first case, total earnings are equal to the sum of wages, $W_{yi}$, plus self-employment income, $I_{yi}$, but are less than the maximum taxable earnings base ($E_{yi} = W_{yi} + I_{yi} < M_y$). In this case, taxable earnings are determined by

$$c)\ T_{yi} = t_y W_{yi} + t_{ys} I_{yi}$$

where $t_{ys}$ = the self-employment OAI tax rate. Case two pertains to the case where total earnings exceed the maximum taxable earnings base, but total wages do not ($E_{yi} > M_y$, but $W_{yi} < M_y$); then,

$$d)\ T_{yi} = t_y W_{yi} + t_{ys}(M_y - W_{yi}).$$

The final case is identical to the pre-1951 formula when taxable earnings, $W_{yi}$, are equal to, less than, or greater than the maximum taxable earnings base. For the years after 1955, total wages are defined as the sum of agricultural and nonagricultural wages and taxable earnings, $T_{yi}$, are calculated using the 1951-1954 formulas.

The revenue stream marked "contributions" qualifies the worker for primary and spousal benefits provided she satisfies the eligibility criteria established by the social security laws effective in the year of retirement. The discounted present value of the expected OAI benefit stream for a single person on the date of retirement is

$$B_i^s = \sum_{K=0}^{99-R} \sum_{t=1}^{12} \frac{R(12) + K(12) + t^{P}R(12)^{b_o}(1+C)^K}{(1+i)^{K(12)+t}} \qquad (4.4)$$

and the discounted present value of a couple's OAI benefit stream is[1]

$$B_i^c = \sum_{K=0}^{99-R} \sum_{t=1}^{12} \frac{b_o(1+C)^K Z}{(1+i)^{K(12)+t}} \qquad (4.5)$$

where

$$Z = R(12) + K(12) + t^{P^M}R(12)^{+} R(12) + K(12) + t^{P^F}R(12) - 0.5\{R(12) + K(12) + t^{P^M}R(12) \cdot R(12) + K(12) + t^{P^F}R(12)\}$$

$t$ = number of benefit payment periods per year,

$99-R$ = number of years in the retirement period,

$R$ = the retirement age of the worker and spouse,

$R(12) + K(12) + t^{P^M}R(12)$ = the probability of the male retiree surviving to life age $R(12) + K(12) + t$ given he is already life age $R(12)$ (expressed in months),

$R(12) + K(12) + t^{P^F}R(12)$ = the probability of the female retiree surviving to life age $R(12) + K(12) + t$ given she is already life age $R(12)$ (expressed in months),

$b_o$ = the initial OAI benefit level received at the end of the first month of retirement,

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[1]For expository convenience, it is assumed in equation (4.4) that the husband and wife are the same age and retire at the same age. In Chapter V of this dissertation, this assumption is dropped.

The Z term captures the joint probability of the household surviving each successive month in the retirement period.

C = the expected growth in prices in subsequent years, and

i = the discount rate.

The life-cycle model of contributions and benefits represented by equations (4.1) through (4.5) captures the salient features of the OAI program. That is, workers pay in a stream of income during their earning years and receive a stream of income in their retirement years, where the right to the benefit stream depends on their past participation on the "contribution" side of the existing program. This is not to imply, however, that the contribution and benefit streams have any tangible relationship except that prior contributions qualify the worker for future benefits. The two revenue streams are not worker specific and need not be comparable in value. The value of the contribution stream depends on the number of earning years, the placement of the earning years in the work history, the worker's taxable earnings in those years, the OAI contribution rate and base, and the interest rate. The value of the benefit stream depends on the discount rate, the growth in future prices, the retiree's life expectancy, and the initial benefit payment. The value of the initial benefit payment, in turn, depends on the worker's average monthly earnings, the progressive benefit formula, age at retirement, familial characteristics, and post-retirement earnings level.

Redistribution, within the intertemporal framework, is determined by the relationship between the total value of the accumulated contributions (4.1, 4.3), and the present discounted value of the expected OAI benefit stream (4.4, 4.5). If the following condition holds for an individual,

$$TC_i^{1,2} \lessgtr B_i^{s,c}, \tag{4.6}$$

then the individual is expected, on average, to receive retirement benefits that are greater than (less than) the accumulated value of her OAI contributions. In this case, the OAI program affects the lifetime income stream for the individual (couple) within the retirement cohort. Similarly, redistribution across cohorts occurs if

$$\sum_{i=1}^{n} TC_i^{1,2} \lessgtr \sum_{i=1}^{n} B_i^{s,c}. \tag{4.7}$$

An actuarially fair retirement would satisfy the following two conditions:

$$TC_i^{1,2} = B_i^{s,c}, \text{ and} \tag{4.8}$$

$$\sum_{i=1}^{n} TC_i^{1,2} = \sum_{i=1}^{n} B_i^{s,c}. \tag{4.9}$$

For instance, if each individual purchases an actuarially fair life annuity with her accumulated contributions, then she can expect, on average, to receive a benefit stream exactly equal to her original lump sum premium (condition 4.8). An annuity purchased with her total OAI contributions at the point of retirement insures the individual against economic risk over an uncertain life span by redistributing income from her relatively high earning years to her low earning years. The value of the monthly annuity payment depends on the value of the lump sum premium, the annuitant's age at retirement, the discount rate, the survivorship table, and the inflation rate (see Chapter V, section F).

Given the above model and definitions of an actuarially fair retirement program, the beneficiary's benefit level can be divided along functional lines. The actuarial component of the individual's OAI benefit payment is the annuity payment which satisfies condition (4.8). The difference between the retiree's 1972 benefit level ($b_o$) and the annuity benefit ($b_a$) would render the amount of redistribution from the program. The redistribution component for individual i is, therefore, defined as follows

$$RC_i = b_{oi} - b_{ai}. \tag{4.10}$$

## V. METHODOLOGY

A life-cycle model of the OAI program is employed to preserve the link between prior OAI contributions paid into the program over the worker's earnings history and OAI benefits received by the beneficiary during retirement. The contributory system modeling of social security is consistent with the individual equity analysis undertaken in this study. However, it is not meant to imply that the contribution and benefit streams have any tangible relationship except that prior contributions "qualify" the worker for future benefits.

The model discussed in the previous section was estimated to examine the impact of differential mortality rates, age at retirement, sex, marital status, income, post-retirement earnings, and price indexing on the OAI redistribution component. In this section, the assumptions of the model, the data set and sorting technique, computational formulas, annuity-type counterfactuals, and redistribution estimates will be briefly discussed.

### A. Fairness Standard

The OAI program can be, and frequently is, evaluated on the basis of two conflicting standards of fairness. If fairness, for instance, is defined as giving more to those persons with a greater relative need, then the social adequacy goal of the program is the main focal point of analysis. The relative need standard of fairness evaluates the program's performance in terms of whether or not greater income protection is extended to those aged persons with greater relative needs, independent of previous OAI contributions. However, if fairness means actuarially fair or, in other words, giving more to those persons with a larger initial investment, then the individual equity goal of the program is emphasized. The relative investment standard of fairness evaluates the performance of the program in terms of actuarially fair rates of return on total OAI contributions. This latter definition of fairness is most frequently used to answer whether or not an individual beneficiary is receiving her (his) "money's worth" from the government program.

In this study, an actuarial standard of fairness is employed to determine what a covered worker would have received from an actuarially fair retirement program.

## B. Study Sample

Data on the socioeconomic characteristics, 1972 OAI benefit level, and OAI benefit and claim status information for persons represented in the study sample were obtained from the 1973 Current Population Survey-Administrative Record Exact Match File. The 1973 Exact Match File unites survey records for persons included in the March, 1973 Current Population Survey to their corresponding benefit and earnings information in the administrative records of the Social Security Administration and to specific items from their 1972 IRS individual income tax returns (Aziz, Kilss, and Scheuren, 1978; Kilss and Scheuren, 1978; Scheuren and Tyler, 1975). Additional earnings information was obtained from the Longitudinal Social Security Exact Match File, 1937-1976. This file includes longitudinal earnings data on adults represented in the 1973 Exact Match File. The study sample included 353 single persons aged 62 and older and 2,771 couples where at least one member was age 62 or older (the data set is described in detail in Appendix B).

A record from the 1973 Exact Match File was included in the study sample if:[1]

1. the individual was 62 or older,
2. retired between 1962 and 1972,
3. represented a "good match,"[2]
4. the claim code in 1972 indicated retired, special age-72 or transitional claim type, and
5. the beneficiary code in 1972 indicated worker only or wife.

This study investigates the OAI program exclusively; hence, reported benefits include primary worker, spousal, transitional, and special age-72 benefits.[3] The level of primary worker benefits received by the

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[1]It was sufficient to have one record in a married couple satisfy the above criteria to get both records included in the sample. Annuity calculations for married persons require the preservation of the family unit.

[2]To be considered a "good match," all members of a stats unit must have matched Summary Earnings Record, Internal Revenue Service and Master Beneficiary Record data present on the file, and a certain level of agreement between demographic information.

[3]Special age-72 benefits are monthly benefits payable to a person aged 72 (before 1972 for male and 1970 for female) or over without sufficient quarters of coverage to qualify for a retired-worker benefit under either the full or transitional insured-status provisions.

Transitional benefits are monthly benefits payable to a person age 72 (before 1969) who has at least one quarter of coverage for every year after 1950 up to the year he/she reached age 65 (male) or 62 (female) with at least three quarters accumulated.

worker beneficiary is a function of the worker's average monthly earnings, age at retirement, and level of post-retirement earnings. Spousal benefits are 50 percent of the retired worker's primary insurance amount adjusted for the spouse's retirement age and post-retirement earnings. All of the aforementioned benefit levels are automatically indexed to the consumer price index beginning in 1975.

OAI beneficiaries are distinguished by the following socioeconomic characteristics:

| | |
|---|---|
| Sex | (male, female), |
| Race | (white, nonwhite), |
| Education | (0-7, 8, 9-11, 12, 13+), |
| Age | (62-64, 65, 66-72, 73+), and |
| Marital status | (married, nonmarried). |

The sex, race, education, and marital status definitions and divisions are consistent with the Kitagawa and Hauser (1973) definitions and divisions. The age divisions are selected to monitor specific features of the social security program.

## C. Historical Contribution Base and Tax Rates

Covered workers and their respective employers are assessed a proportional payroll tax on earnings up to the annual maximum taxable limit. In 1937, a combined employee-employer two percent payroll tax was assessed on the first $3,000. Both the contribution base and payroll tax rate have been periodically increased since 1937. By 1972, the combined tax rate was 9.2 percent applied to the first $9,000. The contribution base and tax rate are based on the historical series located in Appendix C, Table 13.3. The initial impact of the OAI payroll tax rate is shared equally by employees and employers; however, it is assumed that the final burden of the tax is borne by labor, i.e., that there is 100 percent backward shifting.[1]

The historical tax rate series employed in this study is based on the OAI tax rate series constructed by Leimer (1976). Leimer used a historical-net-expenditure-decomposition technique to divide past OASDI contributions along functional lines according to net expenditures on three separate and distinct social security programs: old-age, survivor, and disability. The OAI tax rate series was derived by allocating a share of the OASDI tax rate according to the proportion of total program expenditures accounted for by the OAI portion in every year. Expenditures on old-age insurance differs from survivor and disability insurance

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[1]The shifting assumption is controversial (Brittain, 1971 and 1972a; Feldstein 1972 and 1974; Hammermesh, 1979; MacRae and MacRae, 1976; Munnell, 1974; Vroman, 1974) but conventional in most studies of individual equity (Aaron, 1974; Brittain 1972b; Burkhauser and Warlick, 1981; Freiden, Leimer, and Hoffman, 1976; Leimer, 1978; Okonkwo, 1976; Ozawa, 1974). There are a few computer simulation studies, based on representative individual equity measures, that have attempted to isolate the effect of the shifting assumption in individual equity measures. For instance, Chen and Chu (1974) found that internal rates of return are negatively related and contribution-benefit ratios positively related to the degree of backward shifting, ceteris paribus.

in that OAI represents saving for retirement, whereas SI and DI provide term insurance prior to retirement.[1]

## D. Interest Rates

A low rate of return was selected in determining the compounded value of total OAI contributions. The annual yield on U.S. government bonds from 1937 to 1972 was used in the traditional compounding scheme (see Appendix C, Table 13.1). For the roll-over scheme, the market yield on U.S. government securities at constant maturity from 1937 to 1972 was employed (see Appendix A, Table 11.1). A low rate of return was selected for both compounding schemes because of the riskless nature of the retirement investment. The "absence of risk" is assumed since the government essentially guarantees the worker full repayment of OAI contributions upon retirement.

To further replicate the program's design, the roll-over scheme was introduced into the analysis. The financing design of the social security program is as follows: 1) the government compels covered workers (and their employers) to pay social security taxes; and

[1]The decomposition of OASDI rates is especially important when benefit comparisons are made across women with different labor-homemaker decisions. A working woman covered by social security is eligible for disability benefits and her family is eligible for survivors' benefits, on the basis of her OASDI contributions in the event she should become disabled or die prior to retirement. The nonworking woman and her family are not offered these benefits if the nonworking woman should become disabled or die. The nonworking woman is eligible for disability or survivorship benefits if the disability or death contingency occurs to her husband. Hence, the survivor and disability insurance coverages extended to the working woman prior to retirement are not duplicated by her husbands OASDI contribution.

2) workers do not have access to this money until retirement at which time it is repaid in monthly stipends for life. The OAI contributions are essentially "tied-up" indefinitely. As mentioned earlier, the guaranteed repayment feature implies a riskless investment. The "tied-up" feature suggests a long-term investment, or an investment period equal to the difference between the year of retirement and the year in which the contribution-investment is made. Both the certainty and timing features of the OAI program are reflected in the roll-over scheme.

The roll-over scheme assumes the government invests the worker's contribution into a government bond with the longest maturity that does not exceed the number of years from the date of investment to retirement. Upon maturity, the coupon and principal are immediately rolled over into the next longest bond that has a "correct" maturity period. While it is technically true that the worker could "cash out" of a bond with an "incorrect" maturity period offering a higher yield on the date of retirement, it is assumed that the funds are "tied up" in riskless investments with minimal portfolio management. The difference between the traditional and roll-over total contribution measures are shown in Appendix D, Table 14.1.

## E. Survivor Probabilities

Three tables of survivor probabilities were used to calculate annuity counterfactuals. Survivor probabilities describe the statistical probability of a person life age x (say, 65) surviving to life age x + 1 (say, 66). The age-specific (gender-merged) and age-race-sex-specific tables are based on Social Security Administration (SSA) survivor probabilities for persons 62 and older and Vital Statistics Life Tables for persons younger than 62 (see Appendix C, Tables 13.4 and 13.5). The SSA probabilities were estimated using 1968-1969 Medicare data for persons who were either covered by Hospital Insurance or Supplemental Medical Insurance and at least 62 years old (Bayo, 1972; Myers and Bayo, 1965).

In addition, a table of survivor probabilities differentiated by age, race, sex, marital status, education, and income was used. The socioeconomic adjusted survival probability table is based on tables constructed by Kitagawa and Hauser (1973) and modified by Leimer (1978).

## F. Computational Formulas

Burkhauser and Warlick (1981) estimated a "transfer" component (1972 OASI benefit level less the actuarially fair benefit level) from a life-cycle model using the 1973 Exact Match File. The actuarially fair counterfactual was an immediate whole life annuity payable on an annual basis.[1] This dissertation extends their work to account for the monthly disbursement of benefits and indexing. The annuity is assumed to be purchased on the date of retirement with the retirement candidate's total OAI contributions. The first benefit payment from the actuarially fair retirement insurance is received at the end of the first month of the retirement period.

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[1]A whole life annuity immediate pays the first payment one payment interval after the date of purchase and is purchased with a single premium. See Jordan (1975) for annuity formula derivations.

The variables used to calculate the formulas discussed in this section are as follows:

$PV^S$ = present value of a one dollar unindexed whole life annuity payable monthly,

$PV^C$ = present value of a one dollar unindexed joint-and-two-thirds whole life annuity payable monthly,

$PV^{S*}$ = present value of a one dollar price-indexed whole life annuity payable monthly,

$PV^{C*}$ = present value of a one dollar price-indexed joint-and-two-thirds whole life annuity payable monthly,

$R$ = male's age at retirement,

$\bar{R}$ = female's age at retirement,

${}_{\bar{R}+t}S_{\bar{R}}$ = the probability of the annuitant surviving to life age $\bar{R}+t$ given she is already life age $\bar{R}$,

$101-\bar{R}$ = number of years in the retirement period,

$i$ = discount rate (0.05 percent),

$J$ = wife's age at husband's retirement,

$Z$ = husband's age at wife's retirement,

$s$ = deferment period $|K - Q|$,

$Q$ = retirement age difference between husband and wife $(R - \bar{R})$,

$K$ = age difference between the husband and wife,

$c$ = expected growth in future prices,[1]

$i'$ = indexed discount rate (0.02189 percent), and

[1]Expected growth in future price assumption is based on Trustees intermediate II-B projection on inflation for 1972 of 0.0275.

X = age of the oldest member of the couple at the end of the deferment period.

The retirement candidate purchases an actuarially fair life annuity with her total OAI contributions ($TC_i^{1,2}$) on the date of retirement (RE). The present value of a one dollar nonindexed life annuity payable 12 times a year purchased by a single person is

$$PV^S = \{\sum_{t=1}^{101-\bar{R}} \frac{1}{(1+i)^t}\ {}_{\bar{R}+t}S_{\bar{R}}\} + \frac{11}{24}. \tag{5.1}$$

The present value of a one dollar nonindexed joint-and-two-thirds annuity payable 12 times a year purchased by each member of a couple on the date of retirement is[1]

$$PV^C = [\frac{2}{3}(\underbrace{\{\sum_{t=1}^{101-R} \frac{1}{(1+i)^t}\ {}_{R+t}S_R\} + \frac{11}{24}}_{(a)}) \tag{5.2}$$

$$+ \frac{2}{3}(\underbrace{\frac{1}{(1+i)^s}\ {}_{J+s}S_J}_{(b)}\ \{(\underbrace{\sum_{t=1}^{101-\bar{R}} \frac{1}{(1+i)^t}\ {}_{\bar{R}+t}S_{\bar{R}}) + \frac{11}{24}}_{(c)}\})$$

$$- \frac{1}{3}(\underbrace{\frac{1}{(1+i)^s}\ {}_{J+s}S_J\ {}_{R+s}S_R}_{(d)}\{(\underbrace{\sum_{t=1}^{101-X} \frac{1}{(1+i)^t}\ {}_{\bar{R}+t}S_{\bar{R}} \cdot {}_{Z+t}S_Z) + \frac{11}{24}}_{(e)}\})].$$

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[1]A joint-and-two-thirds is comparable to purchasing a single annuity on each member's life and a joint-and-survivor annuity on both lives. The joint-and-two-thirds replicates the OAI program. The joint-and-two-thirds annuity has an upper bound of one if both members survive, pays 2/3 if there is one survivor, and has a lower bound of zero if there are no survivors in the group.

Term (a) in equation (5.2) is a restatement of equation (5.1) and states that the person purchasing the annuity will receive 2/3 of one if he lives. The second term states that the spouse will receive 2/3 of one when she is eligible for retirement (in the case where the husband is purchasing the annuity). Term (b) is the discount and survivorship factor capturing the deferment period for spousal benefits in the case where the spouse is younger than the husband and not of retirement age. The age difference between the husband and wife is equal to K years, the difference in their retirement ages, $R - \bar{R}$, equals Q, where R is the husband's age at retirement and $\bar{R}$ is the spouse's retirement age. The length of deferment period equals s, where $s = |K-Q|$. If $K = 0$ and $Q = 0$, then $s = 0$ and terms (b) and (c) collapse to $2/3PV^S$—the annuity formula for a single person multiplied by 2/3.

Term (e) in equation (5.2) is a joint-life annuity and it defines group failure when the first member of the group dies or fails to qualify for benefit payments. Failure to qualify in this case means one of the members does not meet the OAI eligibility criteria. The joint-life annuity pays only if both persons are alive <u>and</u> retired and provides payments for the duration of the shorter surviving status. Term (d) accounts for the time value of money and the compound probability of both members surviving the deferment period (s).

The price-indexed annuity formula guarantees payment of a real stream of income over the annuitant's retirement period. The nonindexed formulas discussed above are modified by a CPI expected growth factor and an adjusted interest rate.[1]

The present value of a one dollar price-indexed life annuity payable 12 times a year purchased by a single person is

$$PV^{S*} = \left(\left\{\frac{1}{(1+c)} \sum_{t=1}^{101-\bar{R}} \frac{1}{(1+i')^{t}}\ {}_{\bar{R}+t}S_{\bar{R}}\right\} + \frac{11}{24}\right). \tag{5.3}$$

The present value of a one dollar price-indexed joint-and-two-thirds life annuity payable 12 times per year purchased by each member of a couple is

$$PV^{C*} = \Big[\frac{2}{3}\Big(\Big\{\frac{1}{(1+c)} \sum_{t=1}^{101-R} \frac{1}{(1+i')^{t}}\ {}_{R+t}S_{R}\Big\} + \frac{11}{24}\Big) \tag{5.4}$$

$$+ \frac{2}{3}\Big(\frac{1}{(1+i')^{s}}\ {}_{J+s}S_{J}\ \Big\{\Big(\frac{1}{(1+c)} \sum_{t=1}^{101-\bar{R}} \frac{1}{(1+i')^{t}}\ {}_{\bar{R}+t}S_{\bar{R}}\Big) + \frac{11}{24}\ \Big\}\Big)$$

$$- \frac{1}{3}\ \Big(\frac{1}{(1+i')^{s}}\ {}_{J+s}S_{J}\ {}_{R+s}S_{R}\Big\{\Big(\frac{1}{(1+c)} \sum_{t=1}^{101-X} \frac{1}{(1+i')^{t}}\ {}_{\bar{R}+t}S_{\bar{R}}$$

$$\cdot\ {}_{Z+t}S_{Z}\Big) + \frac{11}{24}\Big\}\Big)\Big].$$

The price index formulas state that the retirement candidate purchases a one dollar life annuity and a series of staggered deferred life annuities paying increments of (1 + c). The nominal accretions in income each year will maintain the real purchasing power of one dollar over the

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[1]The effective interest rate used to calculate an annuity that pays geometrically increasing payments is $i' = \frac{i + c}{1 - c}$, where i is the unindexed interest rate and c is the future growth in prices.

individual's retirement period, assuming that the actual inflation rate equals the expected rate.

## G. Annuity-Type Counterfactuals

There are 12 annuity-type counterfactuals estimated in this study. The counterfactuals are described in Table 5.1. Annuity-type counterfactuals mimic the features of the OAI program and differ in terms of the survivor probability tables used, the compounding scheme employed, and whether benefits are indexed or nonindexed. The value of the monthly annuity benefit is dependent on the accumulated value of OAI contributions, the extent of insurance protection, and the degree to which the insurer can "tailor" benefits to reflect differentials in survivorship.

## H. Earnings Test

The annuity benefits were adjusted for the earnings test. The modeling of the earnings test reflects the legislated earnings test in 1972.

A beneficiary's annuity benefit was adjusted by a reduction factor, $RED_i$, if earnings in 1972 exceeded \$1,680. The reduction factor is calculated by

$$RED_i = 1/2(REP72_i - 1{,}680) \tag{5.5}$$

if $REP72_i \leq 2{,}880$; or

Table 5.1. Description of annuity counterfactuals

| Annuity counterfactual | Characteristics |
|---|---|
| Type 1 | Traditional compounding scheme, nonindexed formula, and gender-merged survivorship tables |
| Type 2 | Traditional compounding scheme, nonindexed formula, and sex-race-distinct survivorship tables |
| Type 3 | Traditional compounding scheme, nonindexed formula, and socioeconomic survivorship tables |
| Type 4 | Traditional compounding scheme, indexed formula, and gender-merged survivorship tables |
| Type 5 | Traditional compounding scheme, indexed formula, and sex-race-distinct survivorship tables |
| Type 6 | Traditional compounding scheme, indexed formula, and socioeconomic survivorship tables |
| Type 7 | Roll-over compounding scheme, nonindexed formula, and gender-merged survivorship tables |
| Type 8 | Roll-over compounding scheme, nonindexed formula, and sex-race-distinct survivorship tables |
| Type 9 | Roll-over compounding scheme, nonindexed formula, and socioeconomic survivorship tables |
| Type 10 | Roll-over compounding scheme, indexed formula, and gender-merged survivorship tables |
| Type 11 | Roll-over compounding scheme, indexed formula, and sex-race-distinct survivorship tables |
| Type 12 | Roll-over compounding scheme, indexed formula, and socioeconomic survivorship tables |

$$RED_i = 600 + (REP72_i - 2{,}880) \qquad (5.6)$$

$$\text{if } REP72_i > 2{,}880$$

where $REP72_i$ = beneficiary i's 1972 reported earnings.

## I. Redistribution Components

The counterfactuals described above were used to calculate the redistribution components, $RC_i$. The redistribution components determine the portion of the beneficiary's 1972 social security benefits which she did not pay for, but which represents an intergenerational transfer from the current working population.

For single beneficiaries, the redistribution components were calculated by the following:

$$RC_{ij} = b_{oi} - b_{ij} \text{ for } j = 1, 2, 3, 7, 8, 9; \text{ and} \qquad (5.7)$$

$$RC^*_{ij} = b_{oi} - b^*_{ij} \text{ for } j = 4, 5, 6, 10, 11, 12 \qquad (5.8)$$

where $RC_{ij}$ and $RC^*_{ij}$ = beneficiary i's redistribution component for annuity-type j,

$b_{oi}$ = 1972 OAI benefit level for beneficiary i,

$b_{ij}$ = nonindexed annuity-type j benefit level for beneficiary i, and

$b^*_{ij}$ = indexed annuity-type j benefit level for beneficiary i.

The redistribution component calculations for married persons are similar to the components calculated for single persons but require the inclusion of both the husband and wife's annuity-type benefit. Family annuity benefits from the joint-and-two-thirds annuity were assumed to be equally owned by the husband and wife. The "equally-owned" assumption has important implications in terms of the relative share of redistribution received by men and women in different household types.

If the husband and wife are retired, then the redistribution components for each member of the couple are calculated by

$$RC_{ij} = b_{oi} - .5(b_{ij} + \underline{b}_{ij}) \text{ for } j = 1, 2, 3, 7, 8, 9; \tag{5.9}$$

$$\underline{RC}_{ij} = \underline{b}_{oi} - .5(b_{ij} + \underline{b}_{ij}) \text{ for } j = 1, 2, 3, 7, 8, 9; \tag{5.10}$$

$$RC^{\dagger}_{ij} = b_{oi} - .5(b^{\dagger}_{ij} + \underline{b}^{\dagger}_{ij}) \text{ for } j = 4, 5, 6, 10, 11, 12; \tag{5.11}$$

$$\underline{RC}^{\dagger}_{ij} = \underline{b}_{oi} - .5(b^{\dagger}_{ij} + \underline{b}^{\dagger}_{ij}) \text{ for } j = 4, 5, 6, 10, 11, 12 \tag{5.12}$$

where $b_{oi}$ = female's 1972 OAI benefit,

$\underline{b}_{o}$ = male's 1972 OAI benefit,

$b_{ij}$ = female i's nonindexed annuity-type j benefit level,

$\underline{b}_{ij}$ = male i's nonindexed annuity-type j benefit,

$b^{\dagger}_{ij}$ = female i's indexed annuity-type j benefit level, and

$\underline{b}^{\dagger}_{ij}$ = male i's indexed annuity-type j benefit level.

If only one member of the couple is retired, then the redistribution calculations are identical to those calculated for single persons.

## J. Behavioral Responses

The removal of the worker-finance retirement insurance was accomplished by estimating a series of worker-specific actuarially fair counterfactuals assuming no behavioral responses. That is, it was assumed that worker participants would not respond by altering their labor or saving decisions when retirement benefits were calculated using strictly-insurance benefit formulas as compared to the OAI retirement benefit formula. An actuarially fair retirement system was used only as a counterfactual to determine the retirement benefits the worker-beneficiary actually paid for through OAI contributions after contributions were already paid into the system. This annuity-type counterfactual was then used to isolate the size of the benefits the beneficiary received from the "social adequacy" function of the government's retirement program. The benefit disentanglement was undertaken with the sole intention of assessing the benefit incidence of the transfers received by the 1972 retirement cohort from the current working population. The incidence was examined to isolate the effects of socioeconomic characteristics on the direction and size of the transfers and to ensure that the intent of the law was consistent with the overall effect of the program.

The <u>ex post</u> annuity calculations and comparisons used in this study are confined to the narrow disentanglement interpretation discussed above. They cannot be accurately interpreted to reflect the effect of a program switch from the current OAI program to an actuarially fair retirement system. Empirical results, to date, show that the social security program does effect labor supply and savings decisions (Boskin, 1977; Burkhauser, 1980; Burkhauser and Quinn, 1981; Feldstein, 1974; Pellechio, 1978). In addition, research by Browning (1975) and Burkhauser and Turner (1978) indicates that an actuarially fair retirement system would have significant labor supply implications across the life cycle.

In light of existing empirical research on the economic effects of the social security program, a study on the privitization of the social security program would necessitate *ex ante* modeling of an actuarially fair retirement system which would fully incorporate behavioral responses by worker participants. At best, this study only approximates the effects of a privitization of the social security program.

# VI. REGRESSION ANALYSIS

In this chapter, the single and married models estimated to isolate the effect of worker characteristics on the percentage of redistribution are presented. Section A includes a description of the generalized quadratic models for single and married households and the model variables. Model variables, independent and dependent, are discussed in detail in Section B.

## A. Functional Form

### 1. Single model

The following generalized quadratic model was estimated to isolate the partial effect of worker-specific characteristics on the percentage of redistribution (X):[1]

$$\begin{aligned} X = {} & \beta_0 + \beta_1 \mathrm{LYEAR} + \beta_2 \mathrm{LYEAR2} + \beta_3 \mathrm{SEX} + \beta_4 \mathrm{RACE} + \beta_5 \mathrm{SERLEN} \\ & + \beta_6 \mathrm{SERLEN2} + \beta_7 \mathrm{RAGER1} + \beta_8 \mathrm{RAGER2} + \beta_9 \mathrm{RAGER3} \\ & + \beta_{10} \mathrm{RCOHORT1} + \beta_{11} \mathrm{RCOHORT2} + \beta_{12} \mathrm{EDU1} + \beta_{13} \mathrm{EDU2} \\ & + \beta_{14} \mathrm{EDU3} + \beta_{15} \mathrm{EDU4} \end{aligned} \tag{6.1}$$

where the dependent and independent variables are defined in Tables 6.1 and 6.2, respectively, and explained in Section VI.B. Four permutations

[1]Loglinear and linear forms were also estimated; however, the quadratic form provided the best fit of the data.

Table 6.1. Definitions of the dependent variables used in the single regression equations

| Variable | Description |
|---|---|
| $RRC_{ij}$ | The nonindexed, nonearnings-adjusted redistribution component for individual j as a percentage of individual j's 1972 OAI benefit level, where i equals type-1, type-2, or type-3 annuity counterfactual. |
| $RRC^*_{ij}$ | The indexed, nonearnings-adjusted redistribution component for individual j as a percentage of individual j's 1972 OAI benefit level, where i equals type-4, type-5, or type-6 annuity counterfactual. |
| $ERRC_{ij}$ | The nonindexed, earnings-adjusted redistribution component for individual j as a percentage of individual j's 1972 OAI benefit level, where i equals type-1, type-2, or type-3 annuity counterfactual adjusted by the earnings test. |
| $ERRC^*_{ij}$ | The indexed, earnings-adjusted redistribution component for individual j as a percentage of individual j's 1972 OAI benefit level, where i equals type-4, type-5, or type-6 annuity counterfactual adjusted by the earnings test. |

Table 6.2. Definitions of the independent variables used in single regression equations

| Variable | Description |
|---|---|
| LTEAR | Accumulated value of lifetime earnings (in hundreds of thousands) |
| LTEAR2 | LTEAR squared |
| SEX | Dummy variable for sex: 0 for male, 1 for female |
| RACE | Dummy variable for race: 0 for white, 1 for nonwhite |
| SERLEN | Service length in covered employment |
| SERLEN2 | SERLEN squared |
| RAGER1 | Dummy variable for retirement age: 1 for age 62-64, 0 otherwise |
| RAGER2 | Dummy variable for retirement age: 1 for age 66-71, 0 otherwise |
| RAGER3 | Dummy variable for retirement age: 1 for age 72 and older, 0 otherwise |
| RCOHORT1 | Dummy variable for retirement cohort: 1 for year 1962-1965, 0 otherwise |
| RCOHORT2 | Dummy variable for retirement cohort: 1 for year 1966-1968, 0 otherwise |
| EDU1 | Dummy variable for years of education completed: 1 for years 0-7, 0 otherwise |
| EDU2 | Dummy variable for years of education completed: 1 for years 9-11, 0 otherwise |
| EDU3 | Dummy variable for years of education completed: 1 for year 12, 0 otherwise |
| EDU4 | Dummy variable for years of education completed: 1 for years 13 or more, 0 otherwise |

of the generalized single model were estimated, where the models differed by specification of the dependent variable only. The purposes of constructing these four different models were, first, to see if variables significant in explaining the percentage of redistribution changed under various counterfactual definitions, and, secondly, to determine if there were any unexpected sign reversals in the parameter estimates. Since this study attempts to account for the effect of worker characteristics and program features on the size of the redistribution component, 12 measures of redistribution were used as dependent variables; each measure was calculated identically, in a technical sense, but different annuity counterfactuals were employed in each measure to net out the "worker-purchased" insurance component. For future reference, the estimation of model 6.1 with dependent variable $RRC_{ij}$, $RRC^*_{ij}$, $ERRC_{ij}$ and $ERRC^*_{ij}$ will be referred to as models 1, 2, 3, and 4, respectively. Each model is estimated using three different mortality rate assumptions.

2. Married model

To isolate the partial effect of family-specific characteristics on the percentage of redistribution for a husband-and-wife family unit (Y), the following generalized quadratic model was estimated:[1,2]

---

[1]Only household units where both the husband and wife were retired in 1972 were included in the data set used to estimate model 6.2.

[2]Loglinear and linear models were also estimated, but the reported model resulted in the best fit of the data.

$$
\begin{aligned}
Y = {} & \beta_0 + \beta_1\text{FLTEAR} + \beta_2\text{FLTEAR2} + \beta_3\text{RACE} + \beta_4\text{SERLEN} \\
& + \beta_{5}\_\text{SERLEN} + \beta_6\text{SERLEN2} + \beta_{7}\_\text{SERLEN2} + \beta_8\text{RAGER1} \\
& + \beta_9\text{RAGER2} + \beta_{10}\text{RAGER3} + \beta_{11}\_\text{RAGER1} + \beta_{12}\_\text{RAGER2} \\
& + \beta_{13}\text{RCOHORT1} + \beta_{14}\text{RCOHORT2} + \beta_{15}\_\text{RCOHORT1} \\
& + \beta_{16}\ \text{RCOHORT2} + \beta_{17}\text{EDU1} + \beta_{18}\text{EDU2} + \beta_{19}\text{EDU3} \\
& + \beta_{20}\text{EDU4} + \beta_{21}\_\text{EDU1} + \beta_{22}\_\text{EDU2} + \beta_{23}\_\text{EDU3} \\
& + \beta_{24}\_\text{EDU4} \qquad (6.2)
\end{aligned}
$$

where the dependent and independent variables are defined in Tables 6.3 and 6.4, respectively, and explained in Section VI.B.

Again, 12 versions of model 6.2 were estimated, differing by dependent variable only. The dependent variables are labelled $FAM_{ij}$, $FAM\#_{ij}$, $EFAM_{ij}$, and $EFAM\#_{ij}$. The estimation of model 6.2 using $FAM_{ij}$, $FAM\#_{ij}$, $EFAM_{ij}$, and $EFAM\#_{ij}$ will be subsequently referred to as models 5, 6, 7, and 8, respectively. Each model is estimated using three different mortality rate assumptions.

## B. Regression Variables

### 1. Dependent variables

Twelve annuity counterfactuals were constructed for each household type, differing by program features or life contingency assumption. Annuity counterfactuals distinguish one dependent variable from another.

Table 6.3. Definitions of the dependent variables used in the married regression equations

| Variable | Description |
| --- | --- |
| $FAM_{ij}$ | The nonindexed, nonearnings-adjusted redistribution component for family j as a percentage of family j's 1972 OAI benefit level, where i equals type-1, type-2, or type-3 annuity counterfactual. |
| $FAM^*_{ij}$ | The indexed, nonearnings-adjusted redistribution component for family j as a percentage of family j's 1972 OAI benefit level, where i equals type-4, type-5, or type-6 annuity counterfactual. |
| $EFAM_{ij}$ | The nonindexed, earnings-adjusted redistribution component for family j as a percentage of family j's 1972 OAI benefit level, where i equals type-1, type-2, or type-3 annuity counterfactual adjusted by the earnings test. |
| $EFAM^*_{ij}$ | The indexed, earnings-adjusted redistribution component for family j as a percentage of family j's 1972 OAI benefit level, where i equals type-4, type-5, or type-6 annuity counterfactual adjusted by the earnings test. |

Table 6.4 Definitions of independent variables used in married regression equations

| Variables | Description |
|---|---|
| FLTEAR | Accumulated value of family lifetime earnings (in hundreds of thousands) |
| FLTEAR2 | FLTEAR squared |
| RACE | Dummy variable for race: 0 for white, 1 for nonwhite |
| SERLEN (_SERLEN) | Service length in covered employment for wife (husband) |
| SERLEN2 (_SERLEN2) | SERLEN (_SERLEN) squared |
| RAGER1 (_RAGER1) | Dummy variable for wife's (husband's) retirement age: 1 for age 62-64, 0 otherwise |
| RAGER2 (_RAGER2) | Dummy variable for wife's (husband's) retirement age: 1 for age 66-71, 0 otherwise |
| RAGER3 | Dummy variable for wife's retirement age: 1 for age 72 and older, 0 otherwise |
| RCOHORT1 (_RCOHORT1) | Dummy variable for wife's (husband's) retirement cohort: 1 for year 1962-1965, 0 otherwise |
| RCOHORT2 (_RCOHORT2) | Dummy variable for wife's (husband's) retirement cohort: 1 for year 1966-1968, 0 otherwise |
| EDU1 (_EDU1) | Dummy variable for years of education completed by wife (husband): 1 for years 0-7, 0 otherwise |
| EDU2 (_EDU2) | Dummy variable for years of education completed by wife (husband): 1 for years 9-11, 0 otherwise |
| EDU3 (_EDU3) | Dummy variable for years of education completed by wife (husband): 1 for year 12, 0 otherwise |
| EDU4 (_EDU4) | Dummy variable for years of education completed by wife (husband): 1 for years 13 or more, 0 otherwise |

Recall, annuity counterfactuals were constructed to disentangle the entitled insurance payment of the OAI benefit from the intergenerational redistribution payment. Counterfactuals range from traditional life annuities based on highly aggregated survivorship assumptions to indexed, earnings adjusted life annuities reflecting highly disaggregated survivorship assumptions. The single and married generalized models are estimated using slightly different dependent variables to isolate how specific survivorship assumptions or program features influence the redistributional incidence of the OAI program. This subsection will describe how each dependent variable was calculated for each household type.

a. <u>Percentage of redistribution for the single model ($RRC_{ij}$, $RRC^*_{ij}$, $ERRC_{ij}$, $ERRC^*_{ij}$)</u> There are four generic measures of redistribution for each single household. Each generic measure is distinguished by a program feature (with or without indexing; with or without earnings adjustments), and, within each measure, three survivorship probabilities assumptions were imposed (gender-merged, sex-race-distinct, socioeconomic-adjusted). The calculations used to determine the percentage of redistribution under various assumptions for single households are as follows:

$$RRC_{ij} = \frac{BEN72_j - TB_{ij}}{BEN72_j} \times 100$$

$$\text{for } i = \text{type-1, type-2, type-3,} \qquad (6.3)$$

$$RRC^*_{ij} = \frac{BEN72_j - TB^*_{ij}}{BEN72_j} \times 100$$

for $i$ = type-4, type-5, type-6, (6.4)

$$ERRC_{ij} = \frac{BEN72_j - ATB_{ij}}{BEN72_j} \times 100$$

for $i$ = type-1, type-2, type-3, (6.5)

$$ERRC^*_{ij} = \frac{BEN72_j - ATB^*_{ij}}{BEN72_j} \times 100$$

for $i$ = type-4, type-5, type-6, (6.6)

where $TB_{ij}$ = nonindexed, nonearnings-adjusted type-$i$ annuity benefit for individual $j$,

$TB^*_{ij}$ = indexed, nonearnings-adjusted type-$i$ annuity benefit for individual $j$,

$ATB_{ij}$ = nonindexed, earnings-adjusted type-$i$ annuity benefit for individual $j$, and

$ATB^*_{ij}$ = indexed, earnings-adjusted type-$i$ annuity benefit for individual $j$.

b. <u>Percentage of redistribution for the married model ($FAM_{ij}$, $FAM^*_{ij}$, $EFAM_{ij}$, $EFAM^*_{ij}$)</u> The four generic measures of redistribution for each married household are:

$$FAM_{ij} = \frac{FBEN72_j - TB_{ij} - \underline{\phantom{x}}TB_{ij}}{FBEN72_j}$$

for $i$ = type-1, type-2, type-3, (6.7)

$$FAM^*_{ij} = \frac{FBEN72_j - TB^*_{ij} - \underline{\phantom{x}}TB^*_{ij}}{FBEN72_j}$$

for $i$ = type-4, type-5, type-6, (6.8)

$$EFAM_{ij} = \frac{FBEN72_j - ATB_{ij} - \_ATB_{ij}}{FBEN72_j}$$

for i = type-1, type-2, type-3, (6.9)

$$EFAM^{*}_{ij} = \frac{FBEN72_j - ATB^{*}_{ij} - \_ATB^{*}_{ij}}{FBEN72_j}$$

for i = type-4, type-5, type-6, (6.10)

where FBEN72 = the sum of the wife and husband's 1972 OAI benefit amounts,

$TB_{ij}(\_TB_{ij})$ = nonindexed, nonearnings-adjusted type-i annuity benefit for the wife (husband) in household j,

$TB^{*}_{ij}(\_TB^{*}_{ij})$ = indexed, nonearnings-adjusted type-i annuity benefit for the wife (husband) in household j,

$ATB_{ij}(\_ATB_{ij})$ = nonindexed, earnings-adjusted type-i annuity benefit for the wife (husband) in household j, and

$ATB^{*}_{ij}(\_ATB^{*}_{ij})$ = indexed, earnings-adjusted type-i annuity benefit for the wife (husband) in household j.

2. Independent variables

a. Accumulated value of lifetime earnings (LTEAR, LTEAR2, FLTEAR, FLTEAR2) The lifetime earnings variables (LTEAR, LTEAR2; FLTEAR, FLTEAR2) are two of four quantitative variables included in the generalized polynomial model. LTEAR reflects the individual's lifetime earnings stream on the date of retirement by a single number. FLTEAR is the sum of the husband and wife's lifetime earnings streams. LTEAR and

FLTEAR were expected to have negative coefficients, whereas LTEAR2 and FLTEAR2 were expected to have positive coefficiente. The summary measure of lifetime earnings was calculated assuming:

(1) annual reported taxable earnings ($REP_i$) were received at the beginning of each year; and

(2) the earnings stream was truncated on the date of retirement (YBEGIN2).

Accordingly, the present value of the worker's lifetime real taxable earnings on the date of retirement is:

$$LTEAR = \sum_{i=EYEAR}^{YBEGIN2} \frac{REP_i}{C_i} \left[ \prod_{j=i}^{YBEGIN2} (1 + r_j) \right] \tag{6.11}$$

where YBEGIN2 = year of retirement,

EYEAR = year of entry into covered employment,

$REP_i$ = annual reported earnings in year i,

$C_i$ = consumer price index in year i, and

$r_j$ = annual real interest rate in year j.

The percentage distribution of LTEAR for single households only appears in Table 6.5. Table 6.6 displays the percentage distribution of FLTEAR for married households.

The summary measure of lifetime earnings differs from the simple sum of annual reported earnings by the weighting of annual reported earnings by the annual real interest rate in each year. This weighting scheme was introduced to approximate the individual's lifetime income status on the date of retirement. The compounding rate was a simple historical average of the yield on U.S. government securities (low yield) and the annual

Table 6.5. Percentage distribution of LTEAR, single population only[a]

| LTEAR[b] | Total | Men | Women |
|---|---|---|---|
| 0 - 19.5 | 13.0 | 10.9 | 14.4 |
| 19.6 - 41.8 | 12.7 | 17.4 | 9.8 |
| 41.9 - 65.2 | 8.5 | 7.2 | 9.3 |
| 65.3 - 86.2 | 5.9 | 8.7 | 4.2 |
| 86.3 - 106.9 | 7.4 | 5.8 | 8.4 |
| 107.0 - 129.5 | 6.8 | 6.5 | 7.0 |
| 129.6 - 150.4 | 5.1 | 2.9 | 6.5 |
| 150.5 - 168.8 | 2.5 | 0.7 | 3.7 |
| 168.9 - 195.5 | 4.8 | 2.9 | 6.0 |
| 195.6 - 217.3 | 6.8 | 8.0 | 6.0 |
| 217.4 - 238.8 | 5.4 | 5.1 | 5.6 |
| 238.9 - 260.7 | 2.5 | 0.7 | 3.7 |
| 260.8 - 281.6 | 4.2 | 5.1 | 3.7 |
| 281.7 - 302.8 | 3.1 | 3.6 | 2.8 |
| 302.9 - 325.1 | 3.7 | 5.8 | 2.3 |
| 325.2 - 345.6 | 2.0 | 1.4 | 2.3 |
| 345.7 - 361.7 | 1.7 | 2.2 | 1.4 |
| 361.8 - 388.3 | 2.0 | 2.9 | 1.4 |
| 388.4 - 401.9 | 0.1 | 0.7 | 0.5 |
| 401.9+ | 1.1 | 1.4 | 0.9 |

[a]Totals may not add to 100 because of rounding.

[b]Reported in thousands.

Table 6.6. Percentage distribution of FLTEAR by household type[a]

| FLTEAR[b] | Total | One Earner | Two Earner |
|---|---|---|---|
| 0 - 40.0 | 9.2 | 14.0 | 3.1 |
| 40.1 - 80.3 | 10.4 | 14.6 | 5.0 |
| 80.4 - 119.8 | 8.9 | 10.8 | 6.5 |
| 119.9 - 160.4 | 7.7 | 7.9 | 7.5 |
| 160.5 - 200.0 | 7.2 | 6.3 | 8.5 |
| 200.1 - 241.1 | 8.2 | 7.8 | 8.6 |
| 241.2 - 281.4 | 7.4 | 7.6 | 7.2 |
| 281.5 - 321.9 | 7.2 | 5.8 | 9.0 |
| 322.0 - 362.2 | 8.5 | 8.3 | 8.6 |
| 362.3 - 402.4 | 8.5 | 9.1 | 7.7 |
| 402.5 - 442.1 | 7.6 | 5.4 | 10.4 |
| 442.2 - 482.3 | 3.2 | 1.4 | 5.4 |
| 482.4 - 522.5 | 2.1 | 0.1 | 3.7 |
| 522.6 - 562.3 | 2.1 | 0.0 | 4.6 |
| 562.4 - 596.8 | 0.7 | 0.0 | 1.5 |
| 596.9 - 637.9 | 0.4 | 0.0 | 1.0 |
| 638.0 - 677.5 | 0.4 | 0.0 | 1.0 |
| 677.6 - 712.8 | 0.3 | 0.0 | 1.0 |
| 712.9 - 730.3 | 0.1 | 0.0 | 0.0 |
| 730.4 - 805.2 | 0.1 | 0.0 | 0.0 |

[a]Totals may not add to 100 because of rounding.

[b]Reported in thousands.

yield on corporate paper plus the rate of increase of average stock prices (high yield). The historical average series was converted to real terms since the annual reported earnings were deflated by the consumer price index (see Appendix C, Table 13.1).

There are obvious problems with the LTEAR measure of lifetime income. First, the selection of an appropriate compounding rate or compounding series is somewhat arbitrary. (The sensitivity of the regression results to the compounding series should be investigated in the future.) Second, LTEAR is based on annual reported earnings to social security only; hence, it systematically excludes nonlabor earnings and labor earnings above the taxable earnings ceiling. The third problem with the LTEAR measure involves the actual size of the annual taxable earnings reported in the file. The size of the annual taxable earnings depends on the tax base and types of occupations covered under the law. These policy variables depend on policy decisions and, as a result, policy decisions influence the size of the calculated lifetime earnings measure. The last problem is technical in nature. Annual reported earnings for 1937 to 1950 were not reported annually; rather, the Longitudinal Exact Match File reported a 14-year summary earnings figure. However, the file reports estimated annual quarters of coverage by year for the 1937 to 1950 time period. The year-specific estimated annual quarters of coverage were used to disaggregate the 1937-1950 summary taxable earnings measure. The disaggregation procedure is described in Appendix F.

In spite of the aforementioned problems with the LTEAR measure of lifetime earnings, it is, in the researcher's opinion, the best measure available given the information on the worker's earnings history obtained from the Longitudinal Earnings Match File. In the context of this study, the most serious shortcoming of the LTEAR measure is the systematic exclusion of nonlabor earnings and earnings above the taxable maximum. The gravity of the problem is challenged, however, by the percentage distribution of the LTEAR shown in Tables 6.5 and 6.6. But, as a safeguard, a graduated education level variable was included in the regression analysis, since, generally speaking, there is a positive, although not perfect, correlation between income and education levels.

b. Socioeconomic variables (SEX, RACE, EDU) The SEX, RACE, and EDU dummy variables represent the expected value of the absolute difference in the dependent variable for each beneficiary characteristic, *ceteris paribus*.

The SEX variable was included to monitor the effect, if any, of sex differentials, be it longevity or employment differences, on the extent of redistribution. The dummy variable takes on a value of one when identifying a female. When the annuity counterfactuals reflect survivorship differentials by sex, the coefficient on SEX was expected to be positive.

The RACE variable reflects the race of the family unit, and it was included to determine if race influenced the size of the redistribution component. RACE equals one for nonwhites and zero for whites. RACE was expected to have a negative coefficient when mortality differentials by race were accounted for in the annuity counterfactual.

The EDU variable was included to supplement the earnings measure (LTEAR, FLTEAR) as discussed earlier, and to account for the independent associations of education level on survivorship. Four education classifications were used: EDU1 for persons with 0-7 years of education: EDU2 for persons with 9-11 years of education; EDU3 for high school graduates; and EDU4 for persons with any college education. The coefficients on the EDU variables measure the differential impact of the indicated category and the category of persons with eight years of schooling (the median years of schooling for this age cohort). The coefficient on EDU1 was expected to be positive without adjusting for education differentials in survivorship, but negative if education differentials were introduced into the annuity counterfactual. Coefficients on EDU3 and EDU4 were expected to have a negative sign without adjusting for education differentials in survivorship and may be positive after adjusting for education differentials in survivorship. The sign of the coefficient for EDU2 may be positive or negative. The sign reversal for the EDU3 and EDU4 was expected because education level is inversely related to mortality; hence, the annuity benefit received by persons with high education levels were lower (therefore, their redistribution components larger), <u>ceteris paribus</u>, when survivorship differentials by education level were used to calculate annuity benefits. Education mortality differentials counteract the progressive features of the benefit formula.

c. <u>Program-worker variables (SERLEN, SERLEN2, RAGER1, RAGER2, RAGER3, RCOHORT1, RCOHORT2)</u> SERLEN, a continuous variable, is a single number representing the number of years of nonzero reported earnings. The summary measure was constructed by counting the number of years from the year of entry into the labor force and the year of retirement when annual reported earnings were nonzero. Since workers with longer earnings history pay in more taxes, SERLEN was expected to have a negative coefficient. The coefficient on SERLEN2 was not predicted.

The $RAGER_i$ and $RCOHORT_i$ dummy variables represent the expected value of the absolute difference in the dependent variable for each program-worker characteristic, <u>ceteris paribus</u>.

RAGER1, RAGER2, and RAGER3 isolate the importance of retirement age of the beneficiary on the size of the intergenerational transfer. The retirement age variable did not appear on the file, but with the use of variables on the file, it was possible to construct it, as follows:

$$RAGE = LAGE - (72-YBEGIN2) \qquad (6.12)$$

where LAGE is the beneficiary's age in 1972, and YBEGIN2 is the year the beneficiary retired. If RAGE equalled 62-64, then a code of one was assigned to RAGER1; if RAGE equalled 66-71, then a code of one was assigned to RAGER2; RAGE greater than 72 was coded as one for RAGER3. The comparison group for this dummy series was persons with a RAGE equal to 65; that is, beneficiaries who began receiving benefits at age 65. Previous empirical evidence suggests that RAGER1, RAGER2, and RAGER3 would have negative coefficients.

The retirement cohort dummy variables RCOHORT1 and RCOHORT2 measure the significance of the year of retirement in explaining the variation in the size of the transfer component. Persons retiring between 1962 and 1964 were in the earliest cohort labelled RCOHORT1. Persons retiring between 1965 and 1968 were in the middle cohort labelled RCOHORT2. The retirement cohorts dated after 1968 were used as the control group. A positive sign was expected on coefficients for RCOHORT1 and RCOHORT2. A positive sign was expected because earlier cohorts benefited from the relative immaturity of the program, which made possible extremely low tax rates and frequent increases in benefit levels.

# VII. RESULTS

## A. Descriptive Statistics

### 1. The benefit incidence of the 1972 old-age insurance program, all households

Table 7.1 displays the estimated benefit incidence of the OAI program in 1972 for the 1962-1972 retirement cohorts based on type-6 annuity counterfactual. In the aggregate, 7.09 million dollars in OAI benefits were paid to retired beneficiaries in this subsample; approximately 89 percent of the benefits received were transfers from the current working generation. The $6.3 million in intergenerational transfers were not, however, evenly distributed across the income groups. Contrary to the "social adequacy" objective, the low-income groups ($0-3,000) represented 15 percent of the sample and they received ten percent of the intergenerational transfers, whereas the middle-income groups ($3,001-8,000) received 57 percent of the transfers but represented 53 percent of the sample. The high-income groups ($8,001 plus) received 33 percent of the transfers, but included 32 percent of the sample (see Appendix E, Table 15.1 for the aggregate figures associated with Table 7.1). In absolute terms, the middle-income groups received the largest share of the intergenerational transfers.

The extent of the intracohort redistribution may be inferred from the absolute and relative size of the redistribution component across family income classes. Column 3 in Table 7.1 indicates that all income groups have received more than their "money's worth" from the social

Table 7.1. Benefit incidence of the 1972 old-age insurance program

| Total family income in 1972[a] | (1) OAI benefit level in 1972 (mean) | (2) Type-6 actuarially fair benefit, earnings adjusted (mean)[b] | Redistribution component: (3) Absolute difference (1)-(2) | Redistribution component: (4) Percentage difference $\frac{(1)-(2)}{(1)} \times 100$ | Household population distribution (in percents) |
|---|---|---|---|---|---|
| $ 0- 1,000 | 698 | 17 | 681 | 97.6 | 1 |
| 1,001- 1,500 | 1,065 | 76 | 989 | 92.9 | 1 |
| 1,501- 2,000 | 1,369 | 119 | 1,250 | 91.3 | 3 |
| 2,001- 2,500 | 1,618 | 141 | 1,477 | 91.3 | 5 |
| 2,501- 3,000 | 1,847 | 173 | 1,674 | 90.6 | 5 |
| 3,001- 3,500 | 2,071 | 220 | 1,851 | 89.4 | 6 |
| 3,501- 4,000 | 2,275 | 258 | 2,017 | 88.7 | 8 |
| 4,001- 5,000 | 2,499 | 287 | 2,212 | 88.5 | 13 |
| 5,001- 6,000 | 2,571 | 312 | 2,259 | 87.9 | 11 |
| 6,001- 8,000 | 2,517 | 312 | 2,205 | 87.6 | 15 |
| 8,001-10,000 | 2,381 | 281 | 2,100 | 88.2 | 9 |
| 10,001-20,000 | 2,271 | 240 | 2,031 | 89.4 | 18 |
| 20,001+ | 2,425 | 260 | 2,165 | 89.3 | 5 |
| Total | $7.09[c] | $.796[c] | $6.294[c] | 88.8 | 3,106 |
| Mean | $2,283 | $256 | $2,027 | 88.8 | |

[a]Total family income includes OAI benefits in 1972.

[b]Annuity counterfactual based on the traditional compounding scheme, an indexed annuity formula and socioeconomic survivorship tables.

[c]In millions of dollars.

security program, since for each income class, the mean OAI benefit level (column 1) is larger than the actuarially fair benefit level (column 2). However, the largest relative gains were realized by low-income families. On average, the lowest income family group received $698 annually from OAI, of which $681, or 97.6 percent, was a result of the "social adequacy" feature of the program. Column 4 shows that the redistribution component, as a percentage of the mean OAI benefit level in 1972, generally decreased as the family income level in 1972 increased. This general pattern would seem to suggest that the progressive benefit formula and minimum benefit provisions effectively redistributed income in favor of lower income households; that is, the program in 1972 was progressive.

There are several approaches that could be used to assess the overall progressivity of the OAI program. One approach is based on end-point comparisons. That is, the percentage of redistribution for the lowest income group is compared to the comparable measure for the highest income group. The relatively small low-to-high differential, 97.6 to 89.3 in column 4, suggests that the redistribution formula in 1972 was "mildly" progressive. Another approach evaluates progressivity in terms of a steadily falling percentage of redistribution as the income level increases. It is interesting to note that the redistribution measure in column 4 of Table 7.1 falls steadily as income rises (with the exception of $2,001-2,500) until the $8,001-10,000 income group, after which the percentage of redistribution generally increases. The falling pattern for nine out of 13 income groups would, again, suggest that the program was "generally" progressive. An alternative approach is to evaluate the program's overall progressivity by comparing the highest income group's percentage of redistribution to the percentage of redistribution for all other family income categories. A "truly" progressive program would have a steadily falling, positive differential as income increases, whereas a "truly" regressive program would have a steadily falling, negative differential as income increases. This type of comparison for the results presented in column 4 is displayed in Figure 7.1, curve 1. Clearly, the OAI program demonstrated "truly" progressive features at income levels less than $3,501, but it displayed regressive, although not "truly" regressive, features at income levels greater than $3,500 but less than $10,001.

The different approaches used to assess progressivity can lead to different program assessments from the same descriptive statistics. The "end-point" approach indicates that the OAI program in 1972 was "mildly" progressive, whereas the "patterned" approach shows it to be "generally" progressive throughout the income classifications. However, the "high-income-group-comparison" approach shows that the program exhibited classic progressive features for low-income groups only, and it exhibited strong regressive features for all other income groups except the penultimate income group. The different approaches when taken separately can result in misleading and "over-optimistic" program performance assessment, but, when taken together, the different approaches render a complete depiction of the program's overall performance. That is, the OAI program in 1972 was "mildly" and "generally" progressive across

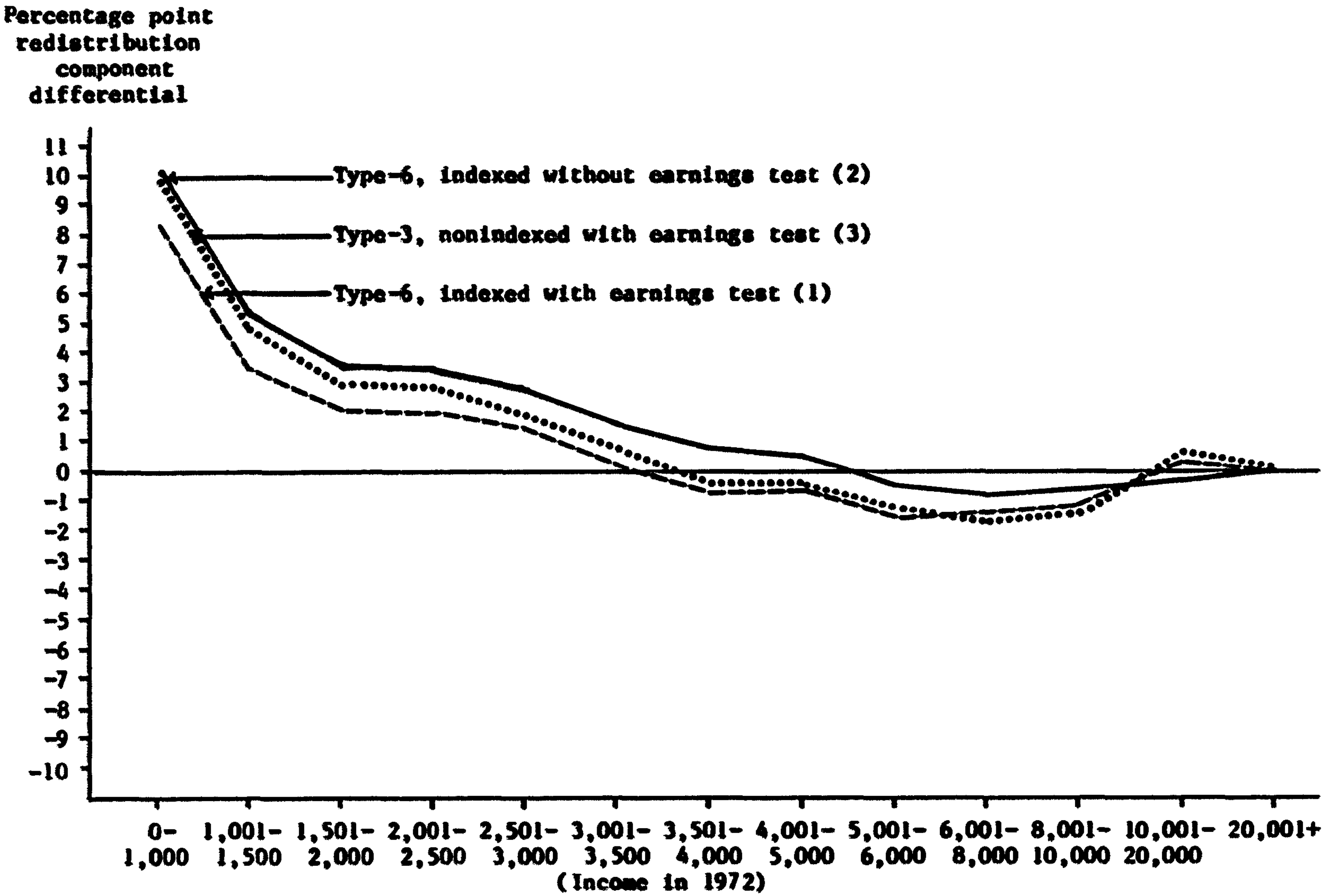


Figure 7.1. Progressivity of the OAI program using socioeconomic-adjusted annuity benefits controlling for earnings test and indexing

income groups, but it also exhibited strong regressive features, resulting in lower relative returns to middle-income beneficiaries. Therefore, the intracohort transfer mechanism operated to pay the highest relative return to the low-income beneficiaries and the lowest relative returns to middle-income beneficiaries, which, in spite of being "mildly" and "generally" progressive, is inconsistent with the program's overall objective.

While the benefit formula and the minimum benefit provisions strongly influenced the pattern of the redistribution components, there are other confounding program features that exert an influence on the redistribution design, such as the earnings test, cost-of-living adjustments, and life contingencies. Table 7.2 isolates the effects of the earnings test and cost-of-living features on the percentage of redistribution across income groups. The life contingency influence is examined in Table 7.3.

The earnings-test effect is presented in column 3 of Table 7.2.[1] Column 3 measures the change in the percentage of redistribution when the earnings test is introduced into the program's design. Note that the earnings test does not affect the three lowest income groups, but it becomes an increasingly important influence on the estimated percentage of redistribution as family income level increases. The earnings-test effect has its greatest impact on high-income families ($6,000+), which

[1]The earnings test operates to reduce the beneficiary's annuity benefit by 50 cents for every dollar of post-retirement earnings greater than $1,680 but less than $2,280 and by $1.00 for every dollar of earnings over $2,280 providing the beneficiary is younger than 72.

Table 7.2. Effect of the earnings test and cost-of-living indexing on the distribution of redistribution (expressed in percentage terms) for socioeconomic-adjusted annuity benefits[a]

| Total family income in 1972[c] | Percentage of redistribution[b] | | | | |
|---|---|---|---|---|---|
| | (1) Indexed and earnings adjusted | (2) Indexed without earnings adjustment | (3) Change in redistribution (1)-(2) | (4) Nonindexed and earnings adjusted | (5) Change in redistribution (1)-(4) |
| $ 0- 1,000 | 97.6 | 97.6 | 0.0 | 97.3 | 0.3 |
| 1,001- 1,500 | 92.9 | 92.9 | 0.0 | 92.5 | 0.4 |
| 1,501- 2,000 | 91.3 | 91.3 | 0.0 | 90.4 | 0.9 |
| 2,001- 2,500 | 91.3 | 91.1 | 0.2 | 90.4 | 0.9 |
| 2,501- 3,000 | 90.6 | 90.3 | 0.3 | 89.5 | 1.1 |
| 3,001- 3,500 | 89.4 | 89.1 | 0.3 | 88.2 | 1.2 |
| 3,501- 4,000 | 88.7 | 88.4 | 0.3 | 87.4 | 1.3 |
| 4,001- 5,000 | 88.5 | 88.2 | 0.3 | 87.0 | 1.5 |
| 5,001- 6,000 | 87.9 | 87.5 | 0.4 | 86.3 | 1.6 |
| 6,001- 8,000 | 87.6 | 86.8 | 0.8 | 85.8 | 1.8 |
| 8,001-10,000 | 88.2 | 87.1 | 1.1 | 86.3 | 1.9 |
| 10,001-20,000 | 89.4 | 87.6 | 1.8 | 87.8 | 1.6 |
| 20,001+ | 89.3 | 87.7 | 1.6 | 87.6 | 1.7 |
| Mean | 88.8 | 88.0 | 0.8 | 87.2 | 1.6 |

[a]All annuity benefits were calculated using socioeconomic-adjusted survivorship probabilities.

[b]Percentage of redistribution was calculated by taking the difference between the mean OAI benefit level in 1972 and the mean actuarially fair benefit level for an income class divided by the mean OAI benefit level in 1972.

[c]Total family income includes OAI benefits received in 1972.

Table 7.3. Changes in the percentage of redistribution under different survivorship probability assumptions

| Total family income in 1972 | Earnings test adjusted: Annuity-type, indexed (1) Type-4[a] | (2) Type-5[a] | (3) Type-6[a] | (4) Change in percentage of redistribution (2)-(1) | (5) (3)-(1) |
|---|---|---|---|---|---|
| $ 0- 1,000 | 97.7 | 97.8 | 97.6 | 0.1 | -0.1 |
| 1,001- 1,500 | 93.9 | 94.1 | 92.9 | 0.2 | -1.0 |
| 1,501- 2,000 | 91.8 | 92.0 | 91.3 | 0.2 | -0.5 |
| 2,001- 2,500 | 91.3 | 91.6 | 91.3 | 0.3 | 0.0 |
| 2,501- 3,000 | 90.6 | 91.0 | 90.6 | 0.4 | 0.0 |
| 3,001- 3,500 | 89.4 | 89.7 | 89.4 | 0.3 | 0.0 |
| 3,501- 4,000 | 88.6 | 88.9 | 88.7 | 0.3 | 0.1 |
| 4,001- 5,000 | 88.2 | 88.6 | 88.5 | 0.4 | 0.3 |
| 5,001- 6,000 | 87.5 | 87.9 | 87.9 | 0.4 | 0.4 |
| 6,001- 8,000 | 87.1 | 87.5 | 87.6 | 0.4 | 0.5 |
| 8,001-10,000 | 87.7 | 88.1 | 88.2 | 0.4 | 0.5 |
| 10,001-20,000 | 88.9 | 89.3 | 89.4 | 0.4 | 0.5 |
| 20,001+ | 88.6 | 88.9 | 89.3 | 0.3 | 0.7 |
| Mean | 88.5 | 88.8 | 88.8 | 0.3 | 0.3 |

[a]Raw data used to calculate the percentage of distribution for each family income classification is available upon request.

places upward pressure on their summary redistribution measures because their annuity benefits are reduced by the earnings test formula. The earnings-test effect on the "high-income-group-comparison" approach to progressivity assessment can be seen by comparing curve 2 to curve 1 in Figure 7.1. In the absence of the earnings test, the program exhibited "classic" progressive features at income levels less than $5,001 and "classic" regressive features at income levels in excess of $5,000. In conclusion, it has been shown that the introduction of the earnings test shifts the performance curve downward, intercalating additional regressive features into the program's _modus operandi_.

Column 5 in Table 7.2 isolates the change in the redistribution measure as a result of introducing price indexing into the program's design. It is interesting to note that the absolute size of the redistribution measure is increased for all income groups when inflation protection is included in the annuity counterfactual, _ceteris paribus_. This result is expected, at least initially, since the indexed annuity benefit is smaller than an unindexed annuity benefit.[1] This is because the annuitant is insured against the risk of economic insecurity _and_ inflation over an uncertain retirement period.

Although all income groups realized extra redistribution per dollar of OAI benefit when indexing was included in the program, the greatest relative gains were realized by higher income groups because of their longer life expectancies on average. Price indexing, when taken alone,

[1]The magnitude of the program-type annuity benefit differential will diminish and its sign will eventually reverse over time because annuity benefits received from an indexed program are augmented by $(1+c)^t$ and unindexed benefits remain fixed in nominal terms.

did not alter the progressivity conclusions, but it did generally reduce the level of progressivity at income levels less than $3,501 and slightly increased regressivity at income levels between $3,500 and $5,000 (see Figure 7.1, curves 3 and 1).

The sensitivity of the progressivity conclusions to the survivorship probability assumption is examined in Table 7.3. The benefit incidence for type-4, type-5, and type-6 counterfactuals are presented in columns 1, 2, and 3, respectively. Column 4 shows the change in the percentage of redistribution if the program adopted a sex-race-age discriminating policy as opposed to strictly age discriminating policy. The adoption of a sex-race-age discriminating policy resulted in an average gain of 0.3 cents of redistribution per dollar of OAI benefit. However, the adoption of a sex-race-age-education-income-marital status discriminating policy in place of an age-only policy (column 5) resulted in a marginal accretion in redistribution for households with income levels in excess of $3,500, where the marginal gain generally increased as family income increased. The lowest income groups ($0-2,000), on the other hand, realized a net loss in redistribution per dollar of OAI benefit. The marginal gain-loss observation is explained by the effect of income and education levels on longevity. That is, annuity benefits are higher (lower) for low (high) income earners, *ceteris paribus*, because the probability of survival is positively related to income and education. Contrary to Aaron's study (1974), the effect of socioeconomic differentials in survivorship does not reverse the direction of redistribution, but, rather, "dampens" the extent of redistribution at the low end of the income scale and "elevates" the extent of redistribution at the high end of the income scale. The program's overall progressivity was virtually invariant to the use of gender-merged or sex-race-distinct survivorship rates (see curves 2 and 3 in Figure 7.2). However, the use of socioeconomic-adjusted survivorship probabilities did augment the regressive features and attentuate the progressive features relative to the "less" discriminating probabilities.

Basic summary statistics for counterfactuals one through six are presented in Tables 7.4 and 7.5. The total and mean annuity benefit received in 1972 and the mean percentage of redistribution, controlling for survivorship assumption, indexing, and earnings test, are presented in Table 7.4. The "end-point" summary statistics for all counterfactuals are shown in Table 7.5. It is interesting to note that the largest progressivity gap (12.2) resulted from a program characterized by age-only discrimination without an earnings penalty test or inflation protection. The smallest progressivity gap (8.3) resulted from a program that provided inflation protection, garnished a fraction of benefits for excessive post-retirement earnings, and tapered benefits to reflect socioeconomic differentials in mortality.

2. <u>The effect of differential life expectancies of males and females on the benefit incidence for fully-insured beneficiaries</u>

a. <u>Single beneficiaries</u> Tables 7.6 and 7.7 show the effect of differential life expectancies of females and males on the benefit incidence for fully-insured single beneficiaries. Type-4 annuity benefits were calculated employing gender-merged survivorship rates,

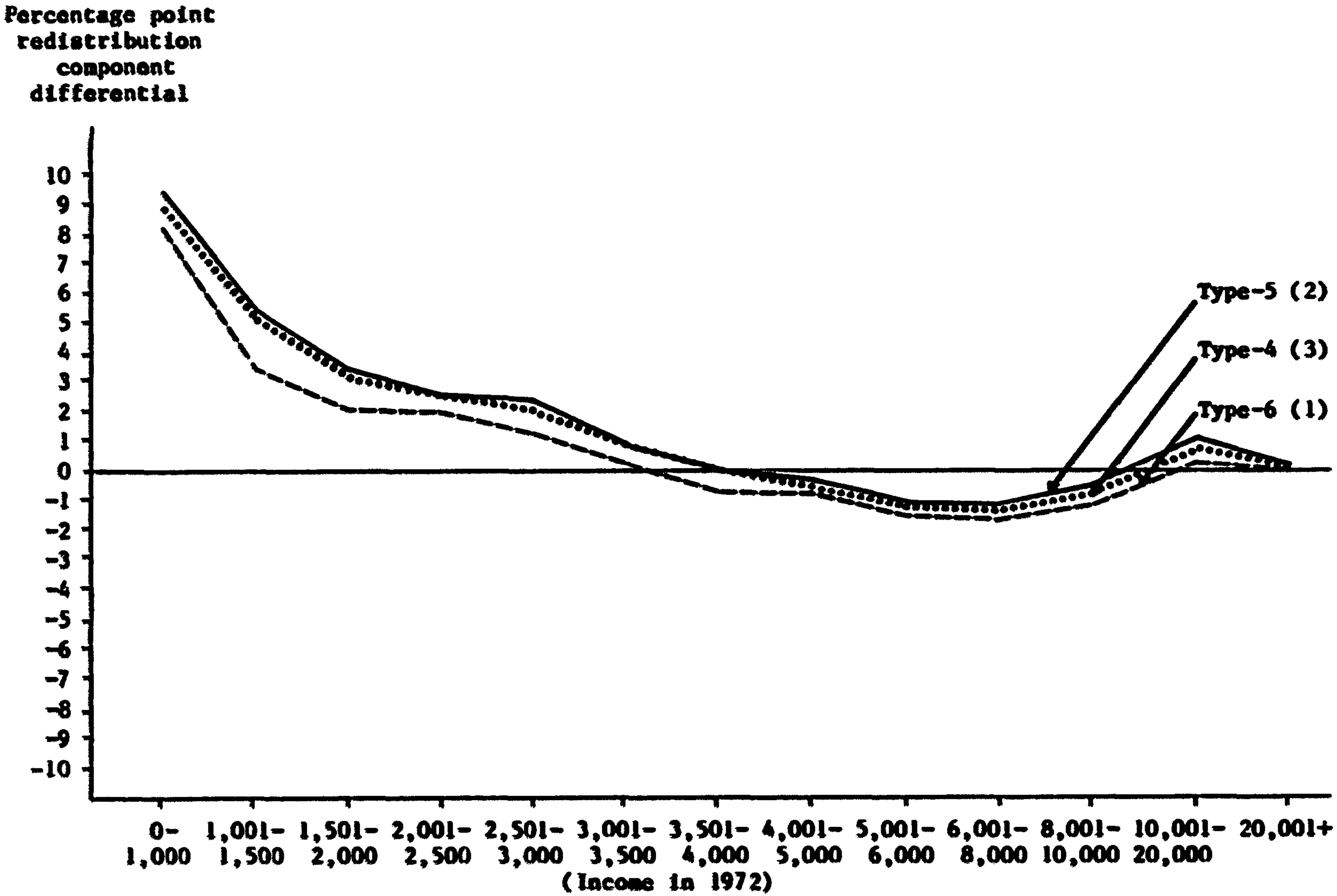


Figure 7.2. Progressivity of the OAI program using different survival probability assumptions

Table 7.4. Total annuity benefit received in 1972 controlling for survivorship assumption, indexing, and earnings test

| Annuity type | Without earnings test: Total annuity benefits | Mean annuity benefit level | Mean[a] percentage of redistribution | With earnings test: Total annuity benefits | Mean annuity benefit level | Mean[a] percentage of redistribution |
|---|---|---|---|---|---|---|
| Unindexed | | | | | | |
| Type-1[b] | $992,000 | 319.00 | 86.0 | $925,000 | 298.00 | 87.0 |
| Type-2[c] | 971,000 | 313.00 | 86.3 | 905,000 | 291.00 | 87.3 |
| Type-3[d] | 972,000 | 313.00 | 86.3 | 906,000 | 292.00 | 87.2 |
| Indexed | | | | | | |
| Type-4[b] | 877,000 | 282.00 | 87.6 | 820,000 | 264.00 | 88.5 |
| Type-5[c] | 850,000 | 274.00 | 88.0 | 794,000 | 256.00 | 88.8 |
| Type-6[d] | 852,000 | 274.00 | 88.0 | 796,000 | 256.00 | 88.8 |

[a]Total benefits minus total annuity benefits divided by total benefits.

[b]Calculations based on gender-merged survivor probabilities.

[c]Calculations based on sex-race-distinct survivor probabilities.

[d]Calculations based on socioeconomic-adjusted survivor probabilities.

Table 7.5. Percentage point gap between poorest and richest income groups

| Survivor probability assumption | Without earnings test | | | | With earnings test | | | |
|---|---|---|---|---|---|---|---|---|
| | Unindexed | | Indexed | | Unindexed | | Indexed | |
| | Poorest to richest | Second poorest to richest | Poorest to richest | Second poorest to richest | Poorest to richest | Second poorest to richest | Poorest to richest | Second poorest to richest |
| Gender-merged | 12.2 | 7.9 | 10.7 | 6.9 | 10.3 | 6.0 | 9.1 | 5.3 |
| Sex-race-distinct | 12.1 | 7.9 | 10.4 | 6.7 | 10.2 | 6.0 | 8.9 | 5.2 |
| Socioeconomic | 11.6 | 6.8 | 9.9 | 5.2 | 9.7 | 4.9 | 8.3 | 3.6 |

Table 7.6. Effect of differential life expectancies of females on benefit incidence for single workers controlling for total OAI contributions[a]

| Total OAI[e] contributions in 1972 dollars | (1) Type-4[b] actu-arially fair benefit | (2) Type-5[c] actu-arially fair benefit | (3) Benefit differ-ential (2)-(1) | (4) Type-5 actu-arially fair benefit | (5) Type-6[d] actu-arially fair benefit | (6) Benefit differ-ential (5)-(4) | (7) Overall benefit differ-ential (5)-(1) | Population |
|---|---|---|---|---|---|---|---|---|
| $ 500< | 22 | 19 | -3 | 19 | 21 | 2 | -1 | 17 |
| 501-1,000 | 62 | 54 | -8 | 54 | 62 | 8 | 0 | 15 |
| 1,001-1,500 | 94 | 81 | -13 | 81 | 87 | 6 | -7 | 13 |
| 1,501-2,000 | 142 | 122 | -20 | 122 | 130 | 8 | -12 | 12 |
| 2,001-2,500 | 183 | 158 | -25 | 158 | 173 | 15 | -10 | 15 |
| 2,501-3,000 | 211 | 182 | -29 | 182 | 198 | 16 | -13 | 15 |
| 3,001-3,500 | 229 | 197 | -32 | 197 | 216 | 19 | -13 | 12 |
| 3,501-4,000 | 293 | 252 | -41 | 252 | 277 | 25 | -16 | 13 |
| 4,001-4,500 | 350 | 301 | -49 | 301 | 320 | 19 | -30 | 17 |
| 4,501-5,000 | 402 | 346 | -56 | 346 | 378 | 32 | -24 | 7 |
| 5,001-6,000 | 410 | 351 | -59 | 351 | 364 | 13 | -46 | 36 |
| 6,001-7,000 | 492 | 421 | -71 | 421 | 456 | 35 | -36 | 12 |
| 7,001-8,000 | 544 | 466 | -78 | 466 | 506 | 40 | -38 | 9 |
| 8,001-9,000 | 767 | 660 | -107 | 660 | 697 | 37 | -70 | 6 |
| 9,001+ | 626 | 537 | -89 | 537 | 562 | 25 | -64 | 8 |
| Total | $59,764 | $51,355 | -$8,409 | $51,355 | $54,853 | $3,498 | -$4,911 | 207 |

[a]Female beneficiaries are defined as single female retirees who are fully insured and collecting primary benefits in 1972.

[b]Type-4 annuity estimates are based on gender-merged survivor probabilities, unadjusted.

[c]Type-5 annuity estimates are based on sex-race-distinct survivor probabilities, unadjusted.

[d]Type-6 annuity estimates are based on socioeconomic adjusted survivor probabilities, unadjusted.

[e]The 1972 dollar value of OAI contributions paid by the worker over her work history. The OAI contributions were accumulated assuming that there was 100 percent backward shifting of the OAI tax rate and compounded at U.S. government bond interest rates.

Table 7.7. Effect of differential life expectancies of males on benefit incidence for single workers controlling for total OAI contributions[a]

| Total OAI[e] contributions in 1972 dollars | (1) Type-4[b] actuarially fair benefit | (2) Type-5[c] actuarially fair benefit | (3) Benefit differential (2)-(1) | (4) Type-5 actuarially fair benefit | (5) Type-6[d] actuarially fair benefit | (6) Benefit differential (5)-(4) | (7) Overall benefit differential (5)-(1) | Population |
|---|---|---|---|---|---|---|---|---|
| $ 500< | 23 | 25 | 2 | 25 | 29 | 4 | 6 | 10 |
| 501-1,000 | 61 | 66 | 5 | 66 | 78 | 12 | 17 | 10 |
| 1,001-1,500 | 91 | 99 | 8 | 99 | 119 | 20 | 28 | 17 |
| 1,501-2,000 | 82 | 88 | 6 | 88 | 104 | 16 | 22 | 4 |
| 2,001-2,500 | 172 | 187 | 15 | 187 | 240 | 53 | 68 | 6 |
| 2,501-3,000 | 215 | 233 | 18 | 233 | 287 | 54 | 72 | 7 |
| 3,001-3,500 | 263 | 284 | 21 | 284 | 347 | 63 | 84 | 11 |
| 3,501-4,000 | 302 | 326 | 24 | 326 | 389 | 63 | 87 | 5 |
| 4,001-4,500 | 313 | 337 | 24 | 337 | 401 | 64 | 88 | 7 |
| 4,501-5,000 | 369 | 399 | 30 | 399 | 484 | 85 | 115 | 7 |
| 5,001-6,000 | 426 | 461 | 35 | 461 | 556 | 95 | 130 | 11 |
| 6,001-7,000 | 469 | 505 | 36 | 505 | 595 | 90 | 126 | 10 |
| 7,001-8,000 | 521 | 561 | 40 | 561 | 664 | 103 | 143 | 11 |
| 8,001-9,000 | 677 | 731 | 54 | 731 | 857 | 126 | 180 | 4 |
| 9,001+ | 827 | 899 | 72 | 899 | 1,100 | 201 | 273 | 6 |
| Total | $37,157 | $40,131 | $2,974 | $40,131 | $48,207 | $8,076 | $11,050 | 126 |

[a]Male beneficiaries are defined as single male retirees who are fully insured and collecting primary benefits in 1972.

[b]Type-4 annuity estimates are based on gender-merged survivor probabilities, unadjusted.

[c]Type-5 annuity estimates are based on sex-race-distinct survivor probabilities, unadjusted.

[d]Type-6 annuity estimates are based on socioeconomic adjusted survivor probabilities, unadjusted.

[e]The 1972 dollar value of OAI contributions paid by the worker over his work history. The OAI contributions were accumulated assuming that there was 100 percent backward shifting of the OAI tax rate and compounded at U.S. government bond interest rates.

whereas type-5 benefits were calculated using sex-race-distinct rates. All contribution classes, independent of sex and annuity type, received positive transfers from the OAI program, i.e., the mean OAI benefit level exceeded the annuity-type benefit level. However, the absolute size of the transfer depends on sex and annuity type. Male beneficiaries received smaller annuity benefits when gender-merged rates were employed relative to a program using sex-distinct rates, ceteris paribus. The observed relationship is expected because sex-distinct rates adjust benefit levels upward for the relatively shorter life expectancies of men, as a group, vis-á-vis women, as a group. Contrariwise, female beneficiaries received larger annuity benefits (hence, smaller redistribution components) when gender-merged rates were used relative to sex-distinct rates. Again, this is an expected result since sex-distinct rates adjust benefit levels downward for the relatively longer life expectancies of women, as a group.

The annuity benefit differentials for female and male beneficiaries are shown in column 3 in Tables 7.6 and 7.7, respectively. The negative differentials for female beneficiaries and the positive differentials for male beneficiaries indicate that single women, as a group, are made differentially better off in a retirement program that does not sex discriminate benefit levels to account for the women's longer life expectancies relative to men's, as a group. Single female beneficiaries, as a group, received annuity benefits that were approximately 16 percent higher in a gender-merged retirement system relative to a sex-race discriminating system, whereas male counterparts, as a group, received benefits that were approximately seven percent lower. Hence, in a sex-neutral retirement program, single male beneficiaries received lower benefit levels relative to a sex discriminating program, which compensated for the slightly higher benefit levels paid to single female beneficiaries.

A similar comparison can be made between type-5 and type-6 annuity counterfactuals. Column 6 in Tables 7.6 and 7.7 shows that single persons, in general, received marginal accretions in their annuity benefits when the effect of marital status, education, and income levels are incorporated into their life contingencies. These "other" socio-economic variables affecting longevity tend to offset the effect of the sex variable for single women and reinforce the effect of the sex variable for single men. The overall benefit differential resulting from the incorporation of sex, race, marital status, education, and income variables into annuity benefit calculations is presented in column 7 on Tables 7.6 and 7.7. Single female beneficiaries received annuity benefits that were approximately eight percent lower in a socioeconomic-discriminating program relative to an age-only discriminating program, whereas single male beneficiaries received annuity benefits that were approximately 30 percent higher.

b. Married beneficiaries Tables 7.8 and 7.9 show the effect of differential life expectancies of females and males on the benefit incidence for fully-insured married beneficiaries. The cross-subsidization by sex found in the case for single beneficiaries was not observed when the annuity benefit comparisons were made across married

Table 7.8. Effect of differential life expectancies of females on benefit incidence for married workers controlling for total OAI contributions, females only

| Total OAI contributions in 1972 dollars[a,b] | (1) Type-4 actuarially fair benefit[c] | (2) Type-5 actuarially fair benefit[d] | (3) Benefit differential (2)-(1) | (4) Type-5 actuarially fair benefit | (5) Type-6 actuarially fair benefit[e] | (6) Benefit differential (5)-(4) | (7) Overall benefit differential (5)-(1) | Population |
|---|---|---|---|---|---|---|---|---|
| $ 500< | 21 | 20 | -1 | 20 | 20 | 0 | -1 | 132 |
| 501-1,000 | 50 | 48 | -2 | 48 | 48 | 0 | -2 | 130 |
| 1,001-1,500 | 83 | 80 | -3 | 80 | 80 | 0 | -3 | 97 |
| 1,501-2,000 | 110 | 107 | -3 | 107 | 105 | -2 | -5 | 89 |
| 2,001-2,500 | 145 | 141 | -4 | 141 | 140 | -1 | -5 | 82 |
| 2,501-3,000 | 172 | 166 | -6 | 166 | 165 | -1 | -7 | 59 |
| 3,001-3,500 | 204 | 198 | -6 | 198 | 194 | -4 | -10 | 51 |
| 3,501-4,000 | 242 | 234 | -8 | 234 | 231 | -3 | -11 | 45 |
| 4,001-4,500 | 275 | 268 | -7 | 268 | 263 | -5 | -12 | 41 |
| 4,501-5,000 | 341 | 329 | -12 | 329 | 324 | -5 | -17 | 23 |
| 5,001-6,000 | 358 | 346 | -12 | 346 | 341 | -5 | -17 | 40 |
| 6,001-7,000 | 360 | 348 | -12 | 348 | 338 | -10 | -22 | 26 |
| 7,001-8,000 | 452 | 441 | -11 | 441 | 435 | -6 | -17 | 17 |
| 8,001-9,000 | 449 | 436 | -13 | 436 | 426 | -10 | -23 | 10 |
| 9,001+ | 479 | 458 | -21 | 458 | 441 | -17 | -38 | 5 |
| Total | $127,919 | $123,835 | -$4,084 | $123,835 | $122,094 | -$1,741 | -$5,825 | 847 |

[a]Female beneficiaries are defined as married female retirees who are fully insured and collecting primary benefits in 1972.

[b]The 1972 dollar value of OAI contributions paid by the worker over her work history. The OAI contributions were accumulated assuming that there was 100 percent backward shifting of the OAI tax rate and compounded at U.S. government bond interest rates.

[c]Type-4 annuity estimates are based on gender-merged survivor probabilities, unadjusted.

[d]Type-5 annuity estimates are based on sex-race-distinct survivor probabilities, unadjusted.

[e]Type-6 annuity estimates are based on socioeconomic adjusted survivor probabilities, unadjusted.

Table 7.9. Effect of differential life expectancies of males on benefit incidence for married workers controlling for total OAI contributions, males only

| Total OAI contributions in 1972 dollars[a,b] | (1) Type-4 actu-arially fair benefit[c] | (2) Type-5 actu-arially fair benefit[d] | (3) Benefit differ-ential (2)-(1) | (4) Type-5 actu-arially fair benefit | (5) Type-6 actu-arially fair benefit[e] | (6) Benefit differ-ential (5)-(4) | (7) Overall benefit differ-ential (5)-(1) | Population |
|---|---|---|---|---|---|---|---|---|
| $ 500< | 20 | 19 | -1 | 19 | 18 | -1 | -2 | 104 |
| 501-1,000 | 46 | 45 | -1 | 45 | 45 | 0 | -1 | 156 |
| 1,001-1,500 | 78 | 75 | -3 | 75 | 75 | 0 | -3 | 134 |
| 1,501-2,000 | 109 | 106 | -3 | 106 | 105 | -1 | -4 | 133 |
| 2,001-2,500 | 130 | 126 | -4 | 126 | 125 | -1 | -5 | 129 |
| 2,501-3,000 | 173 | 168 | -5 | 168 | 166 | -2 | -7 | 138 |
| 3,001-3,500 | 209 | 203 | -6 | 203 | 200 | -3 | -9 | 125 |
| 3,501-4,000 | 234 | 227 | -7 | 227 | 226 | -1 | -8 | 143 |
| 4,001-4,500 | 264 | 256 | -8 | 256 | 254 | -2 | -10 | 133 |
| 4,501-5,000 | 289 | 280 | -9 | 280 | 277 | -3 | -12 | 131 |
| 5,001-6,000 | 330 | 321 | -9 | 321 | 318 | -3 | -12 | 229 |
| 6,001-7,000 | 406 | 395 | -11 | 395 | 389 | -6 | -17 | 214 |
| 7,001-8,000 | 433 | 421 | -12 | 421 | 416 | -5 | -17 | 171 |
| 8,001-9,000 | 478 | 465 | -13 | 465 | 460 | -5 | -18 | 176 |
| 9,001+ | 483 | 469 | -14 | 469 | 462 | -7 | -21 | 247 |
| Total | $647,180 | $629,418 | -$17,762 | $629,418 | $621,542 | -$7,876 | -$25,638 | 2,363 |

[a]Male beneficiaries are defined as married male retirees who are fully insured and collecting primary benefits in 1972.

[b]The 1972 dollar value of OAI contributions paid by the worker over his work history. The OAI contributions were accumulated assuming that there was 100 percent backward shifting of the OAI tax rate and compounded at U.S. government bond interest rates.

[c]Type-4 annuity estimates are based on gender-merged survivor probabilities, unadjusted.

[d]Type-5 annuity estimates are based on sex-race-distinct survivor probabilities, unadjusted.

[e]Type-6 annuity estimates are based on socioeconomic adjusted survivor probabilities, unadjusted.

persons. Actuarially fair benefit levels for married persons were approximately three percent higher, independent of the sex of the primary annuitant, in a retirement system that did not sex discriminate relative to a sex discriminating program (see column 3 of Tables 7.8 and 7.9). First, it is interesting to note that both the male and female received annuity benefits that were three percent higher in a sex-neutral retirement program. Within a married household, the effects of sex differentials are neutralized because the joint-and-two-thirds annuity insures the male and female members of the couple. The absolute size of the annuity benefit received is invariant to the sex of the annuitant who actually purchases the annuity in either program type. Second, the sex-neutral bias in favor of married persons, as a group, is a result of the joint-and-two-thirds annuity, which insures the life of the shorter-lived (on average) male, the longer-lived female, and the longest-lived survivor, who is typically the female. The surviving wife will, in a sex-neutral system, receive artificially high benefit levels for the duration of widowhood. The relatively higher benefit levels for married households in a sex-neutral actuarially fair retirement program are financed primarily by single, male beneficiaries who receive smaller annuity benefits because of the assumption of identical life contingencies for males and females.

The effect of incorporating "other" socioeconomic variables can be seen in column 6 of Tables 7.8 and 7.9. Education, income, and marital status effects tend to further reduce the size of the annuity benefit received by married persons. Specifically, annuity benefits are approximately 1.3 percent lower in a socioeconomic discriminating program relative to a sex-race discriminating program. Again, this is expected since married persons tend to have a longer life expectancy relative to nonmarried counterparts. The overall benefit differential is represented in column 7 on Tables 7.8 and 7.9. Generally speaking, married persons, independent of sex, received benefits that were approximately four percent lower in a socioeconomic discriminating program relative to an age-only discriminating program.

3. <u>The effect of retirement year on the benefit incidence of single workers only</u>

The effect of retirement year on the percentage of redistribution is shown in Table 7.10. The retirement year is divided into three categories: 1962-1965, 1966-1969, and 1970-1972. The results are shown for type-3 and type-6 annuity counterfactuals, and displayed by total family income classifications. Except in a few cases (notably when the cell size is small), the percentage of redistribution falls as the retirement year increases, holding family income constant. Also, the percentage of redistribution is quite stable for the lowest income group, which is consistent with the minimum benefit provision. The generally observed inverse relationship between the percentage of redistribution and the date of retirement supports the findings of Parsons and Munro (1977), Freiden et al. (1976), and Burkhauser and Warlick (1981). The general decline in the redistribution measure reflects the maturing of the program.

Table 7.10. Effect of retirement year on benefit incidence for single workers

| Total family income in 1972 | Type-3, unindexed[a] Redistribution component[b] | | | Type-6, indexed[a] Redistribution component | | | Population | | |
|---|---|---|---|---|---|---|---|---|---|
| | 1962-1965 | 1966-1969 | 1970-1972 | 1962-1965 | 1966-1969 | 1970-1972 | 1962-1965 | 1966-1969 | 1970-1972 |
| $ 500- 1,000 | 97 | 98 | 95 | 97 | 98 | 96 | 1 | 3 | 1 |
| 1,001- 1,500 | 97 | 93 | 76 | 97 | 93 | 83 | 7 | 8 | 6 |
| 1,501- 2,000 | 89 | 84 | 78 | 88 | 85 | 79 | 6 | 9 | 5 |
| 2,001- 2,500 | 88 | 84 | 75 | 87 | 85 | 79 | 18 | 15 | 7 |
| 2,501- 3,000 | 89 | 82 | 79 | 89 | 83 | 82 | 13 | 10 | 9 |
| 3,001- 3,500 | 87 | 81 | 72 | 85 | 82 | 75 | 7 | 13 | 6 |
| 3,501- 4,000 | 87 | 80 | 71 | 87 | 82 | 75 | 12 | 9 | 10 |
| 4,001- 5,000 | 89 | 82 | 73 | 89 | 83 | 76 | 12 | 13 | 12 |
| 5,001- 6,000 | 90 | 79 | 66 | 90 | 80 | 71 | 5 | 12 | 6 |
| 6,001- 8,000 | 91 | 82 | 67 | 90 | 83 | 73 | 6 | 15 | 13 |
| 8,001-10,000 | 87 | 86 | 73 | 86 | 87 | 77 | 5 | 13 | 5 |
| 10,001-20,000 | 94 | 84 | 74 | 94 | 85 | 78 | 16 | 19 | 18 |
| 20,001+ | 88 | 64 | 75 | 88 | 67 | 79 | 4 | 1 | 3 |
| Overall | 90 | 83 | 75 | 90 | 84 | 79 | 112 | 140 | 101 |

[a]Annuity benefits employed to calculate the redistribution components were adjusted for earnings in excess of the 1972 earnings limit.

[b]Redistribution components were calculated by subtracting the mean annuity benefit level from the mean 1972 OAI benefit level reported as a percentage of the mean 1972 OAI benefit level. Raw data used to calculate the reported results are available upon request.

4. <u>The benefit incidence of the 1972 old-age insurance program: Married, both retired households only</u>

a. <u>The effect of the wife's work status on the benefit incidence</u>

There are 1,394 households included in this sample: 614 two-earner households and 780 one-earner households. See Table 7.11 for a description of the married, both retired, data set. The effect of the wife's work status on the distributional impact of the OAI program is examined in Tables 7.12 and 7.13. Female beneficiaries were classified by their work status, where work status was determined by OAI beneficiary eligibility criteria, and household income in 1972. Table 7.12 is similar to Table 7.1 except that only married households where both the husband and wife are retired in 1972 were included in the data set.

Similar to the results in Table 7.1, all female beneficiaries, independent of work status and family income level, received positive income transfers from the OAI program in 1972 (that is, the redistribution components in columns 4a and 8a in Table 7.12 are positive). The redistribution component expressed as a percentage of the female's OAI benefit level is, on average, negatively related to family income, indicative of the program's progressivity.

Table 7.13 compares the differences in OAI benefit level (column 1), yearly annuity benefit in a type-6 actuarially fair retirement system based on the actual contributions made by the female (column 2) and the male (column 3), and redistribution component in percentage terms (column 4) for working and nonworking women across family income categories. The working woman who qualifies for benefits on her own account received, on average, retirement benefits that were approximately

Table 7.11. Population distribution for married, both retired households by family income in 1977 and family type

| Family income in 1972 | Two-earner[a] Population size | Two-earner[a] Percentage distribution | One-earner[b] Population size | One-earner[b] Percentage distribution |
|---|---|---|---|---|
| $ 0- 2,000 | 3 | .5 | 29 | 4.0 |
| 2,001- 2,500 | 13 | 2.0 | 33 | 4.0 |
| 2,501- 3,000 | 13 | 2.0 | 46 | 6.0 |
| 3,001- 3,500 | 31 | 5.0 | 55 | 7.0 |
| 3,501- 4,000 | 46 | 7.5 | 72 | 9.0 |
| 4,001- 5,000 | 107 | 17.5 | 113 | 15.0 |
| 5,001- 6,000 | 91 | 15.0 | 98 | 13.0 |
| 6,001- 8,000 | 122 | 20.0 | 111 | 14.0 |
| 8,001-10,000 | 70 | 11.5 | 66 | 8.0 |
| 10,001-20,000 | 86 | 14.0 | 124 | 16.0 |
| 20,001+ | 32 | 5.0 | 33 | 4.0 |
| Total | 614 | 100.0 | 780 | 100.0 |

[a]Husband and wife are eligible for primary-worker benefits on their own accounts.

[b]Husband is eligible for primary-worker benefits on his own account and the wife is eligible for dependent spouse's benefits only.

Table 7.12. Effect of the wife's work status on wife-only benefit incidence holding family income constant (type-6, earnings adjusted)

| Total family income in 1972[c] | Two-earner household[a] (1) Female OAI benefit level (mean) | (2) Actuarially fair benefit from wife's annuity | (3) Actuarially fair benefit from husband's annuity | (4) Redistribution component (a) 1-2-3 | (4) (b) $\frac{1-2-3}{1}$ | One-earner household[b] (5) Female OAI benefit level (mean) | (6) Actuarially fair benefit from wife's annuity | (7) Actuarially fair benefit from husband's annuity | (8) Redistribution component (a) 5-6-7 | (8) (b) $\frac{5-6-7}{5}$ |
|---|---|---|---|---|---|---|---|---|---|---|
| $ 0- 2,000 | 506 | 29 | 26 | 451 | 89 | 418 | 0 | 35 | 383 | 92 |
| 2,001- 2,500 | 752 | 18 | 60 | 674 | 90 | 542 | 1 | 42 | 499 | 92 |
| 2,501- 3,000 | 1,023 | 38 | 70 | 915 | 89 | 702 | 1 | 89 | 612 | 87 |
| 3,001- 3,500 | 1,193 | 58 | 105 | 1,030 | 86 | 754 | 1 | 102 | 651 | 86 |
| 3,501- 4,000 | 1,210 | 62 | 140 | 1,008 | 83 | 912 | 1 | 143 | 768 | 84 |
| 4,001- 5,000 | 1,255 | 65 | 157 | 1,033 | 82 | 918 | 1 | 138 | 779 | 85 |
| 5,001- 6,000 | 1,316 | 61 | 182 | 1,073 | 82 | 929 | 1 | 172 | 756 | 81 |
| 6,001- 8,000 | 1,413 | 80 | 171 | 1,162 | 82 | 955 | 1 | 162 | 792 | 83 |
| 8,001-10,000 | 1,412 | 88 | 148 | 1,176 | 83 | 896 | 2 | 159 | 735 | 82 |
| 10,001-20,000 | 1,508 | 80 | 129 | 1,299 | 86 | 933 | 1 | 144 | 788 | 84 |
| 20,001+ | 1,596 | 87 | 159 | 1,350 | 85 | 942 | 2 | 134 | 806 | 86 |

[a]Husband and wife are eligible for primary-worker benefits on their own account.

[b]Husband is eligible for primary-worker benefits on his own account and the wife is eligible for dependent spouse's benefits only.

[c]Total family income includes OAI benefits in 1972.

Table 7.13. Comparison of OAI and type-6, earnings adjusted annuity benefits for married women with different labor-homemaker choices holding family income constant

| Total family[e] income in 1972 | (1) Difference[c] between female OAI benefit levels | (2) Difference[b] between actuarially fair benefits from wife's annuity | (3) Difference[a] between actuarially fair benefits from husband's annuity | (4) Difference in[d] redistribution components as a percentage of OAI |
|---|---|---|---|---|
| $ 0- 2,000 | 88 | 29 | -9 | -3 |
| 2,001- 2,500 | 210 | 17 | 18 | -2 |
| 2,501- 3,000 | 321 | 37 | -19 | +2 |
| 3,001- 3,500 | 439 | 57 | 3 | 0 |
| 3,501- 4,000 | 298 | 61 | -3 | -1 |
| 4,001- 5,000 | 337 | 64 | 19 | -3 |
| 5,001- 6,000 | 387 | 60 | 10 | +1 |
| 6,001- 8,000 | 458 | 79 | 9 | -1 |
| 8,001-10,000 | 516 | 86 | -11 | +1 |
| 10,001-20,000 | 575 | 79 | -15 | +2 |
| 20,001+ | 654 | 85 | 25 | -1 |

[a]Fifty percent of the two-earner woman's share of her husband's yearly annuity benefit less 50 percent of the one-earner woman's share of her husband's yearly annuity benefit.

[b]Fifty percent of the two-earner woman's yearly annuity benefit minus 50 percent of the one-earner woman's yearly annuity benefit.

[c]The mean level of OAI benefits received by a woman in a two-earner household less the mean level of benefits received by a woman in a one-earner household.

[d]The difference between redistribution components of women in two-earner and one-earner households.

[e]Total family income includes OAI benefits in 1972.

50 percent larger than the auxiliary benefits received by the nonworking woman. The benefit differential ranges from 21 percent for the lowest income category to 69 percent for the highest income category.[1]

Generally speaking, entitled female workers received retirement benefits that were larger than dependent spouse benefits. One reason for the observed OAI benefit differential is that the nonworking woman's benefit is based on 50 percent of her husband's primary insurance amount, whereas the entitled female worker's benefit is based on her primary insurance amount if her PIA exceeds 50 percent of her spouse's PIA.

Working women received higher annuity benefits from an actuarially fair retirement system based on their actual contributions than nonworking women (column 2, Table 7.13). Column 3 presents the difference between annuity benefits received by working and nonworking women based on actual contributions made by their husbands. The negative values in column 3 indicate that the working woman received a smaller annuity benefit from her husband's joint-and-two-thirds annuity than the nonworking woman. On net, working women received higher annuity benefits based on the household's OAI contributions, and, because of her past contributions, she was afforded higher OAI benefits.

The difference in percentage of redistribution per dollar of OAI benefits for working and nonworking women is shown in column 4 of Table 7.13. Working women received a higher percentage of redistribution in the following income categories: $2,501-3,000, $5,001-6,000,

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[1]Social Security Bulletin data show that the average benefit for women workers to be about 60 percent higher than the wife's auxiliary benefit for this time period.

$8,001-10,000, and $10,001-20,000. But, nonworking women received an equal or higher percentage of redistribution per dollar of OAI benefits in all other income categories. It appears that there was slightly more redistribution to nonworking women <u>vis-à-vis</u> working women. In absolute terms, however, working women paid in more dollars in the form of OAI contributions, and, in exchange, they received higher OAI benefit levels. The relatively narrow differential in redistribution components suggests that women, independent of work status, were treated almost equally in terms of redistribution.

b. <u>The effect of the wife's work status on husband-only benefit incidence</u> The finding of equal treatment across women with different labor-homemaker choices does not apply to men married to women with different labor-homemaker choices. Tables 7.14 and 7.15 represent the male versions of Tables 7.12 and 7.13. It is interesting to note that the male redistribution components as a percentage of OAI benefits (columns 4b and 8b) are generally higher for males in one-earner households relative to their male counterparts in two-earner households. The percentage of redistribution measures follow the generally observed pattern--falling as family income rises. However, the variance in the pattern is slightly smaller for males in a one-earner household (97 to 92 percent). This implies that males in one-earner households with family income of $0-2,000 received 97 cents of redistribution for every dollar of OAI benefit. Similarly, males in the $5,001-10,000 income classes received 92 cents of redistribution per dollar of OAI benefit.

Table 7.14. Effect of the wife's work status on husband-only benefit incidence holding family income constant (type-6, earnings adjusted)

| | Two-earner household[a] | | | | | One-earner household[b] | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) Redistribution component | | (5) | (6) | (7) | (8) Redistribution component | |
| Total family[c] income in 1972 | Male OAI benefit level (mean) | Actuarially fair benefit from wife's annuity | Actuarially fair benefit from husband's annuity | (a) 1-2-3 | (b) $\frac{1-2-3}{1}$ | Male OAI benefit level (mean) | Actuarially fair benefit from wife's annuity | Actuarially fair benefit from husband's annuity | (a) 5-6-7 | (b) $\frac{5-6-7}{5}$ |
| $ 0- 2,000 | 906 | 29 | 26 | 851 | 94 | 1,026 | 0 | 35 | 991 | 97 |
| 2,001- 2,500 | 1,349 | 18 | 60 | 1,271 | 94 | 1,309 | 1 | 42 | 1,266 | 97 |
| 2,501- 3,000 | 1,521 | 38 | 70 | 1,413 | 93 | 1,604 | 1 | 89 | 1,514 | 94 |
| 3,001- 3,500 | 1,598 | 58 | 105 | 1,435 | 90 | 1,727 | 1 | 102 | 1,624 | 94 |
| 3,501- 4,000 | 1,961 | 62 | 140 | 1,759 | 90 | 1,995 | 1 | 143 | 1,851 | 93 |
| 4,001- 5,000 | 1,986 | 65 | 157 | 1,764 | 89 | 2,103 | 1 | 138 | 1,964 | 93 |
| 5,001- 6,000 | 2,150 | 61 | 182 | 1,907 | 89 | 2,092 | 1 | 172 | 1,919 | 92 |
| 6,001- 8,000 | 2,056 | 80 | 171 | 1,805 | 88 | 2,091 | 1 | 162 | 1,928 | 92 |
| 8,001-10,000 | 1,947 | 88 | 148 | 1,711 | 88 | 2,086 | 2 | 159 | 1,925 | 92 |
| 10,001-20,000 | 1,907 | 80 | 129 | 1,698 | 89 | 2,062 | 1 | 144 | 1,917 | 93 |
| 20,001+ | 2,197 | 87 | 159 | 1,951 | 89 | 2,110 | 2 | 134 | 1,974 | 94 |

[a]Husband and wife are eligible for primary-worker benefits on their own account.

[b]Husband is eligible for primary-worker benefits on his own account and the wife is eligible for dependent spouse's benefits only.

[c]Total family income includes OAI benefits in 1972.

Table 7.15. Comparison of OAI and type-6, adjusted annuity benefits for married men in one-earner and two-earner households holding family income constant

| Total family[e] income in 1972 | (1) Difference[c] between male OAI benefit levels | (2) Difference[b] between actuarially fair benefits from wife's annuity | (3) Difference[a] between actuarially fair benefits from husband's annuity | (4) Difference in[d] redistribution components as a percentage of OAI |
|---|---|---|---|---|
| $ 0- 2,000 | -120 | 29 | -9 | -3 |
| 2,001- 2,500 | 40 | 17 | 18 | -3 |
| 2,501- 3,000 | -83 | 37 | -19 | -1 |
| 3,001- 3,500 | -129 | 57 | 3 | -4 |
| 3,501- 4,000 | -34 | 61 | -3 | -3 |
| 4,001- 5,000 | -117 | 64 | 19 | -4 |
| 5,001- 6,000 | 58 | 60 | 10 | -3 |
| 6,001- 8,000 | -35 | 79 | 9 | -4 |
| 8,001-10,000 | -139 | 86 | -11 | -4 |
| 10,001-20,000 | -155 | 79 | -15 | -4 |
| 20,001+ | 87 | 85 | 25 | -5 |

[a]Fifty percent of the two-earner man's yearly annuity benefit minus 50 percent of the one-earner man's yearly annuity benefit.

[b]Fifty percent of the two-earner man's share of his wife's yearly annuity benefit less 50 percent of the one-earner man's share of his wife's yearly annuity benefit.

[c]The mean level of OAI benefits received by a man in a two-earner household less the mean level of benefits received by a man in a one-earner household.

[d]The difference between redistribution components of men in two-earner and one-earner households.

[e]Total family income includes OAI benefits in 1972.

Generally, males in two-earner households received smaller OAI benefits (column 1, Table 7.15), although males in two-earner families received higher combined annuity benefits based on the actual OAI contributions of both earners in the households. The difference in combined annuity benefits (columns 2 plus 3 in Table 7.15) across household type is, in large part, a result of the annuity benefits received from the wife's joint-and-two-thirds annuity based on her actual OAI contributions. Column 4 in Table 7.15 shows that the male in a one-earner household consistently received a larger percentage of redistribution from the OAI program than the male in a two-earner household.

c. <u>The effect of the wife's work status on family benefit incidence</u> Table 7.16 represents the benefit incidence across one-earner and two-earner households, holding constant family income in 1972. Column 7 indicates that, except for the lowest income category, family OAI benefit levels were higher for two-earner households <u>vis-á-vis</u> one-earner households. In addition, two-earner households received higher family benefits from an actuarially fair retirement system (column 8). All family units, independent of household types, received positive income transfers from the OAI program (columns 3 and 6). Furthermore, the one-earner household received a larger percentage of redistribution relative to the two-earner household for all income categories (column 9).

d. <u>The importance of the household type in explaining the benefit incidence</u> The tabular results regarding the percentage of redistribution by sex and household type across family income classes (columns 4b

Table 7.16. Effect of the wife's work status on family benefit incidence holding total family income constant

| | Two-earner household | | | One-earner household | | | Comparison | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Total family income in 1972 | Family OAI benefit level[a] | Family annuity benefit[b] | Redistribution component as a percentage of family OAI benefits (1)-(2) | Family OAI benefit level[a] | Family annuity benefit[b] | Redistribution component as a percentage of family OAI benefits (4)-(5) | Difference in family OAI benefits (1)-(4) | Difference in family annuity benefit (2)-(5) | Difference in redistribution components (3)-(6) |
| $ 0- 2,000 | 1,411 | 111 | 92 | 1,433 | 71 | 95 | -22 | 40 | -3 |
| 2,001- 2,500 | 2,101 | 156 | 93 | 1,851 | 87 | 95 | 250 | 69 | -2 |
| 2,501- 3,000 | 2,544 | 217 | 91 | 2,306 | 181 | 92 | 238 | 36 | -1 |
| 3,001- 3,500 | 2,791 | 328 | 88 | 2,481 | 208 | 92 | 310 | 120 | -4 |
| 3,501- 4,000 | 3,171 | 404 | 87 | 2,907 | 289 | 90 | 264 | 115 | -3 |
| 4,001- 5,000 | 3,242 | 443 | 86 | 3,021 | 279 | 91 | 221 | 164 | -5 |
| 5,001- 6,000 | 3,466 | 487 | 86 | 3,021 | 346 | 89 | 445 | 141 | -3 |
| 6,001- 8,000 | 3,470 | 502 | 86 | 3,046 | 326 | 89 | 424 | 176 | -3 |
| 8,001-10,000 | 3,359 | 472 | 86 | 2,983 | 322 | 89 | 376 | 150 | -3 |
| 10,001-20,000 | 3,414 | 417 | 88 | 2,995 | 289 | 90 | 419 | 128 | -2 |
| 20,001+ | 3,793 | 493 | 87 | 3,051 | 271 | 91 | 742 | 222 | -4 |

[a]Combined OAI benefit received by the husband and wife in 1972.

[b]Combined annuity benefit received by the husband and wife.

and 8b, Tables 7.12 and 7.14 are summarized in Figure 7.3. It is interesting to note that the percentage of redistribution received by women, independant of work status, is generally lower than the comparable measure for men. The observed male-to-female differential in redistribution is consistent across all income categories. But, looking at the redistribution curves for women by household types in Figure 7.3, it appears that the size and pattern of the redistribution measure for women in one-earner and two-earner households are very similar. The observed similarity suggests that, although women with different work statuses paid in different amounts of OAI contributions, they were treated equally in terms of the percentage of OAI benefits representing redistribution from the current working generation.

The redistribution pattern for males in one-earner and two-earner households are similar; however, the absolute size of the redistribution measure varies significantly by household type. It is clear from Figure 7.3 that the percentage of redistribution for males in one-earner households is substantially larger than the comparable measure for males in two-earner households across <u>all</u> income categories. One reason for the obvious size disparity across all income categories is the very small (or zero) annuity benefits received from the nonworking wife's joint-and-two-thirds annuity. Because his wife's yearly annuity value is generally equal to zero, his redistribution component is larger.

Although males in one-earner households received preferential treatment from the OAI program <u>vis-á-vis</u> males in two-earner households

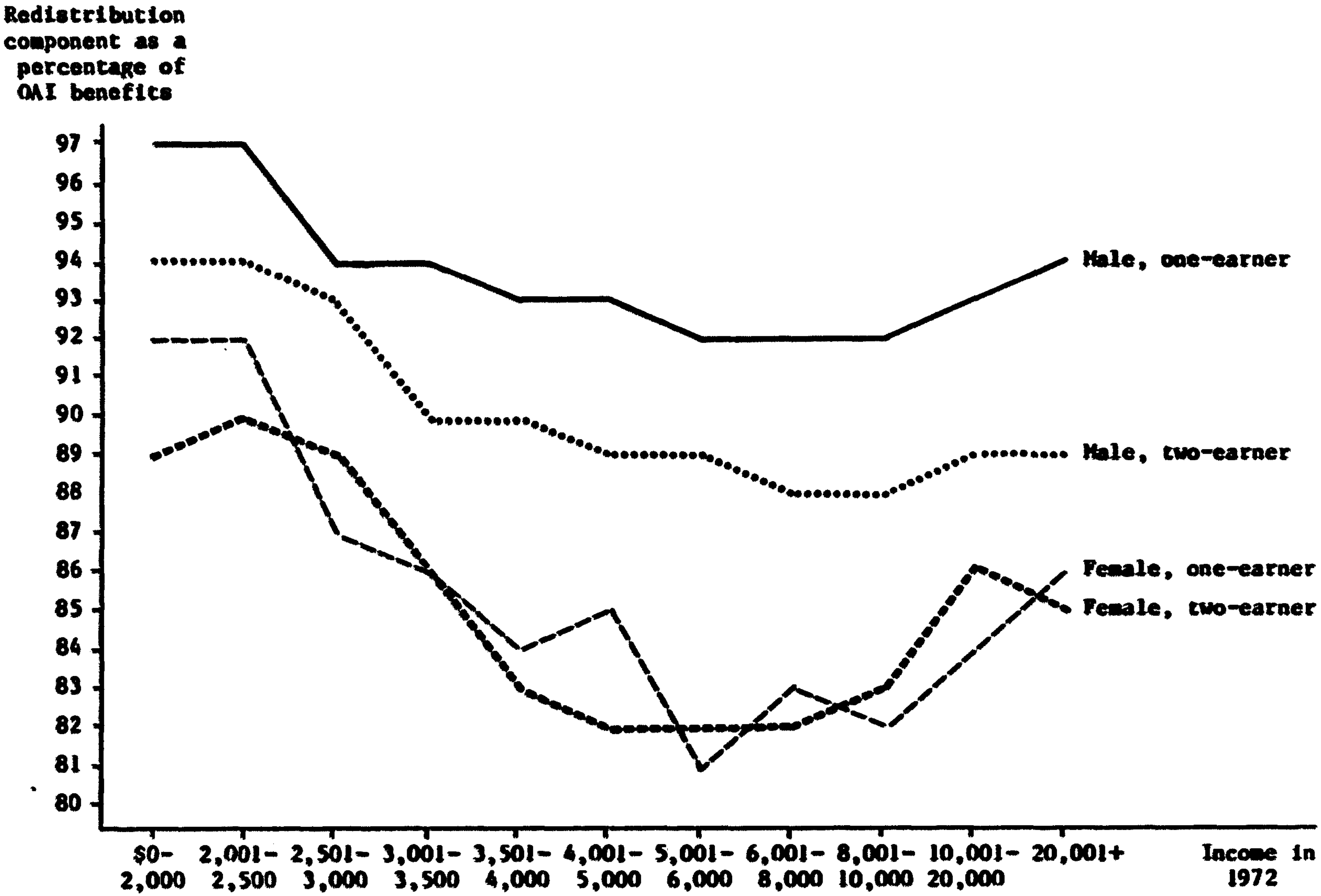



Figure 7.3. Graphical comparison of redistribution components, expressed in percentage terms, by sex and household type

and females, working women, as a group, received a significantly smaller percentage of redistribution when compared to working males. There are several reasons for the smaller redistribution components received by working women. First, entitled women frequently claim reduced benefits. In 1967, 67 percent of the married female retired workers aged 65 and older received reduced benefits. By 1971, the proportion had increased to 76 percent. The proportion of beneficiary women with reduced benefits puts downward pressure on mean OAI benefit levels used to calculate the redistribution components. Second, working women have smaller primary insurance amounts relative to working men because of their lower earnings and intermittent labor force participation. In 1971, a significant proportion of retired women workers, especially the dually entitled, were entitled to the minimum PIA. Half of the dually entitled women workers, in 1971, were entitled to the minimum PIA compared to seven percent of male workers. Differences in PIA distributions for male and female workers reflect differences in work histories. Men generally work for longer periods of time at higher earnings, resulting in higher PIAs. The last reason concerns the annuity benefit received by working women from their husband's past OAI contributions. Since the male worker pays into the system longer and, in addition, receives higher earnings, he has a larger accumulated tax contribution to purchase a joint-and-two-thirds at retirement. Assuming a community property approach to the actuarially fair benefit, the wife receives half of the yearly annuity benefit in an actuarially fair system based on OAI contributions of her husband. The wife's redistribution component is determined by subtracting her OAI benefit level from her share of the yearly family annuity benefit based on her OAI contribution and her husband's OAI contributions. The value of her redistribution component is relatively small, therefore, because her OAI benefit level is generally small because of her smaller PIA relative to male workers combined with her increased tendency to accept reduced benefits and the relatively large annuity benefit received from her husband's joint-and-two-thirds combined with the annuity benefit based on her own OAI contributions.

Figure 7.4 summarizes the tabular results in columns 3 and 6 in Table 7.16. The distribution of redistribution components by household type illustrated in Figure 7.4 shows that one-earner families, on average, received preferential treatment from the OAI program. Again, the preferential status of one-earner families is explained by the nominal contributions made by the nonworking spouse in the one-earner family.

e. <u>The progressivity of the OAI program by household type</u> The "end-point" approach to determining progressivity suggests that the program is "weakly" progressive: the variances for women and men in two-earner families are 90-82 and 94-88, respectively, and the variances for women and men in one-earner families are 92-81 and 97-92, respectively. Progressivity assessment based on the "patterned" approach shows the program to be "generally" progressive given the generally observed inverse relationship between the percentage of redistribution and total family income. However, the "highest-income-group-comparison" approach exposes strong regressive features for women in both household types,

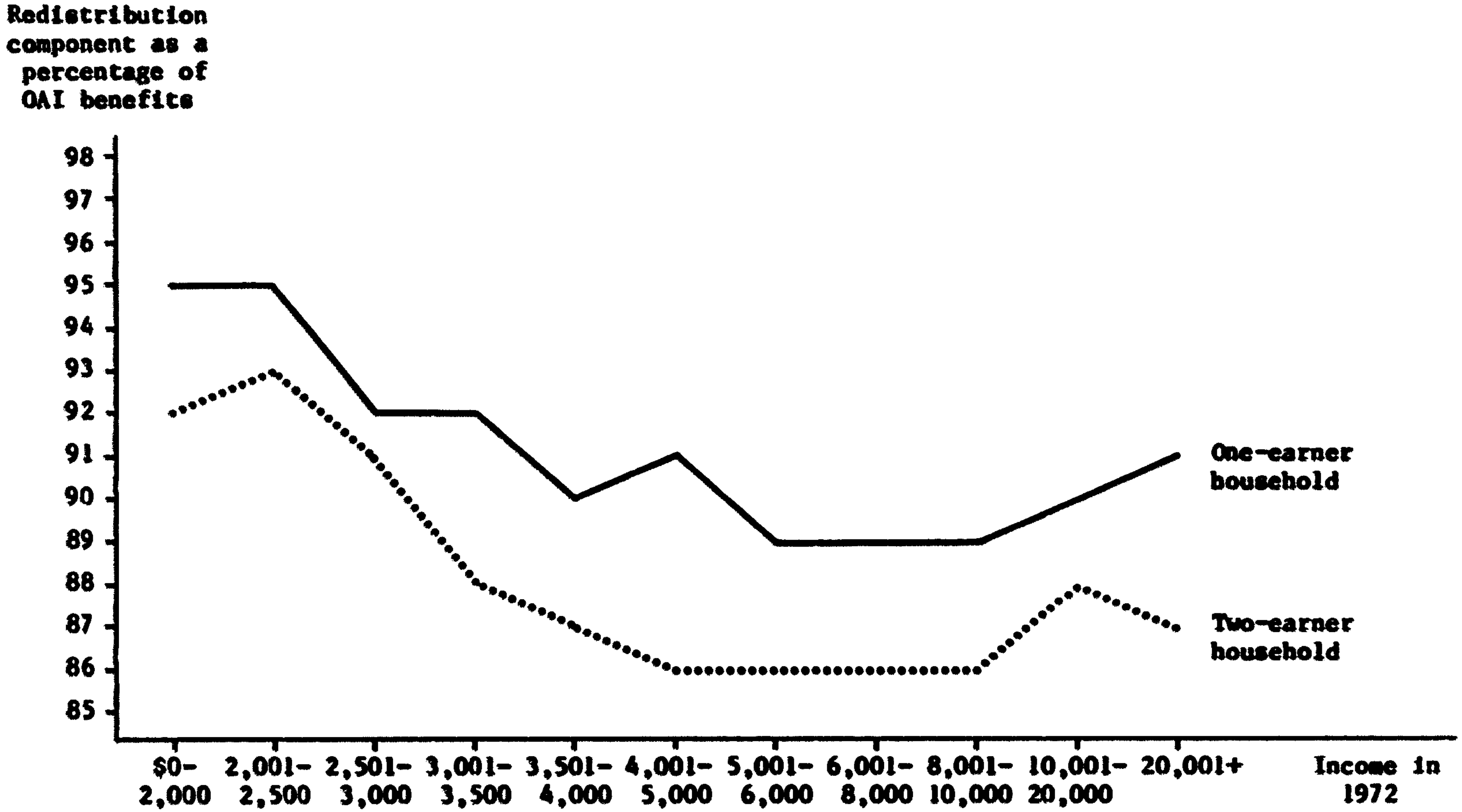
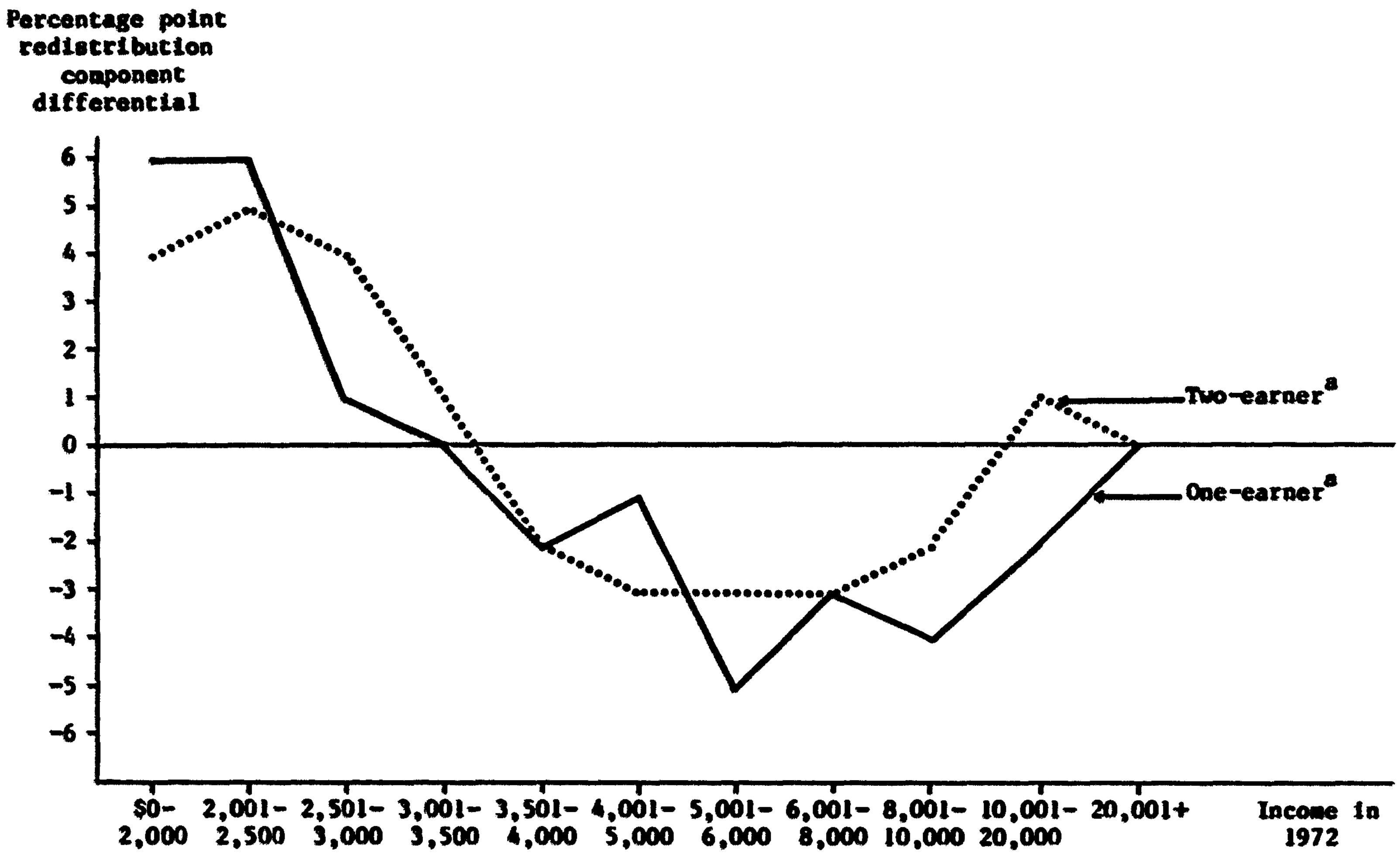



Figure 7.4. Distribution of redistribution components in percentage terms by household type

mild regressive features for men in one-earner households, and strong progressive features for men in two-earner households.

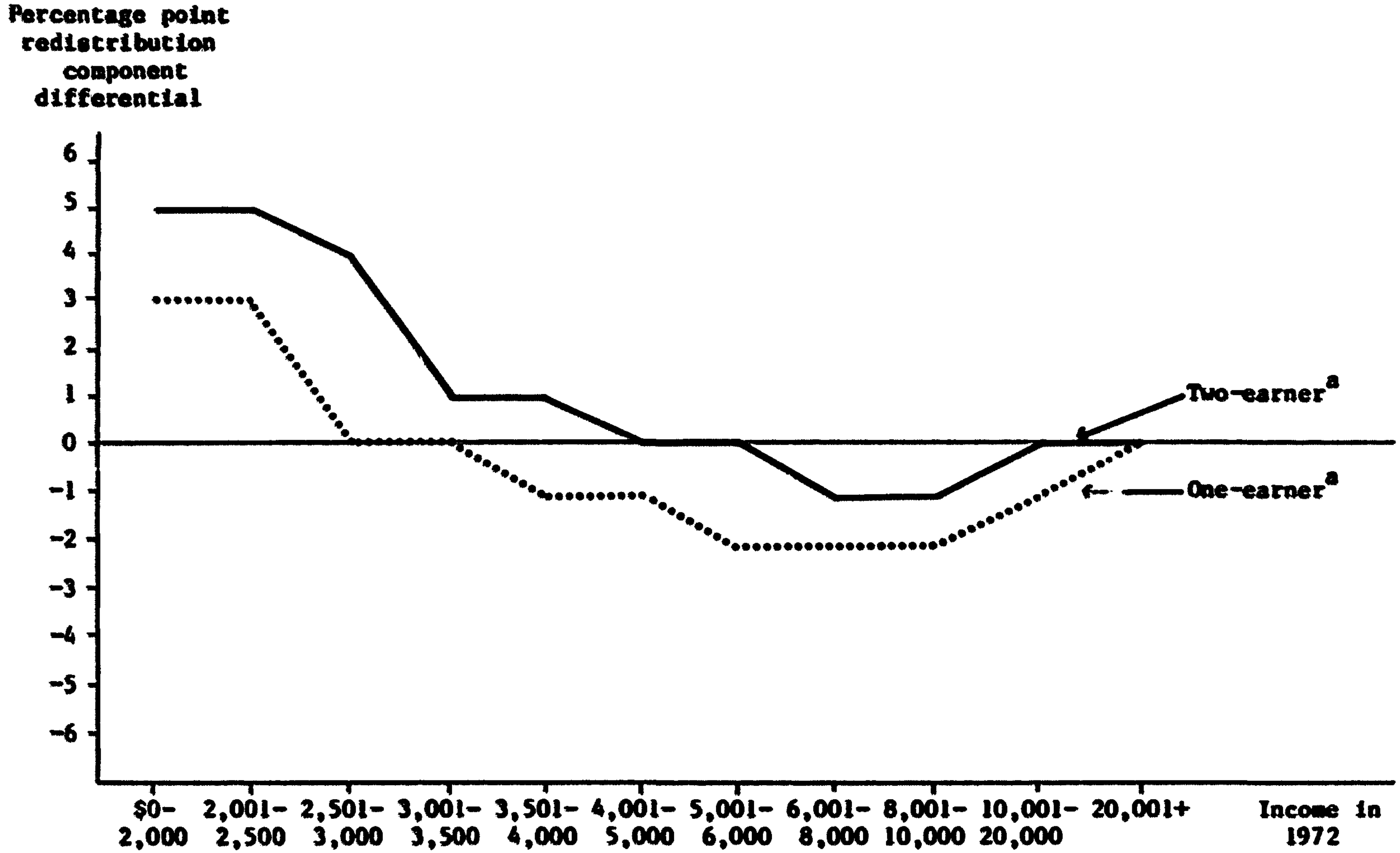
Figure 7.5 applies to females only and shows the program to be progressive at income levels less than $3,500, but strongly regressive at income levels greater than $3,500. Middle-income females, especially, are made worse-off relative to the highest income group of females, independent of household type. The program does not appear to be as regressive when focusing on males only (Figure 7.6). The program demonstrated "classic" progressive features for males in two-earner households for income levels of $5,000 or less, and it demonstrated only "slight" regressive features for the $6,001 to $20,000 range. The program has a narrow progressive area ($0 to $3,000) for males in one-earner households and somewhat "classic" regressive features for income levels in excess of $3,000. Figure 7.7 is based on the household unit sorted by household type. Again, the program had "classic" progressive features at low income levels ($0-3,500), but had regressive features at higher income levels. The program is more progressive and less regressive for two-earner relative to one-earner households.

(Summary findings on annuity types 1, 2, 3, 4, and 5, male-to-female comparisons by annuity type, household-type comparisons by annuity type, and indexed to nonindexed comparisons by annuity type can be found in Appendix E, Tables 15.2, 15.3, 15.4, 15.5, 15.6, and 15.7.)



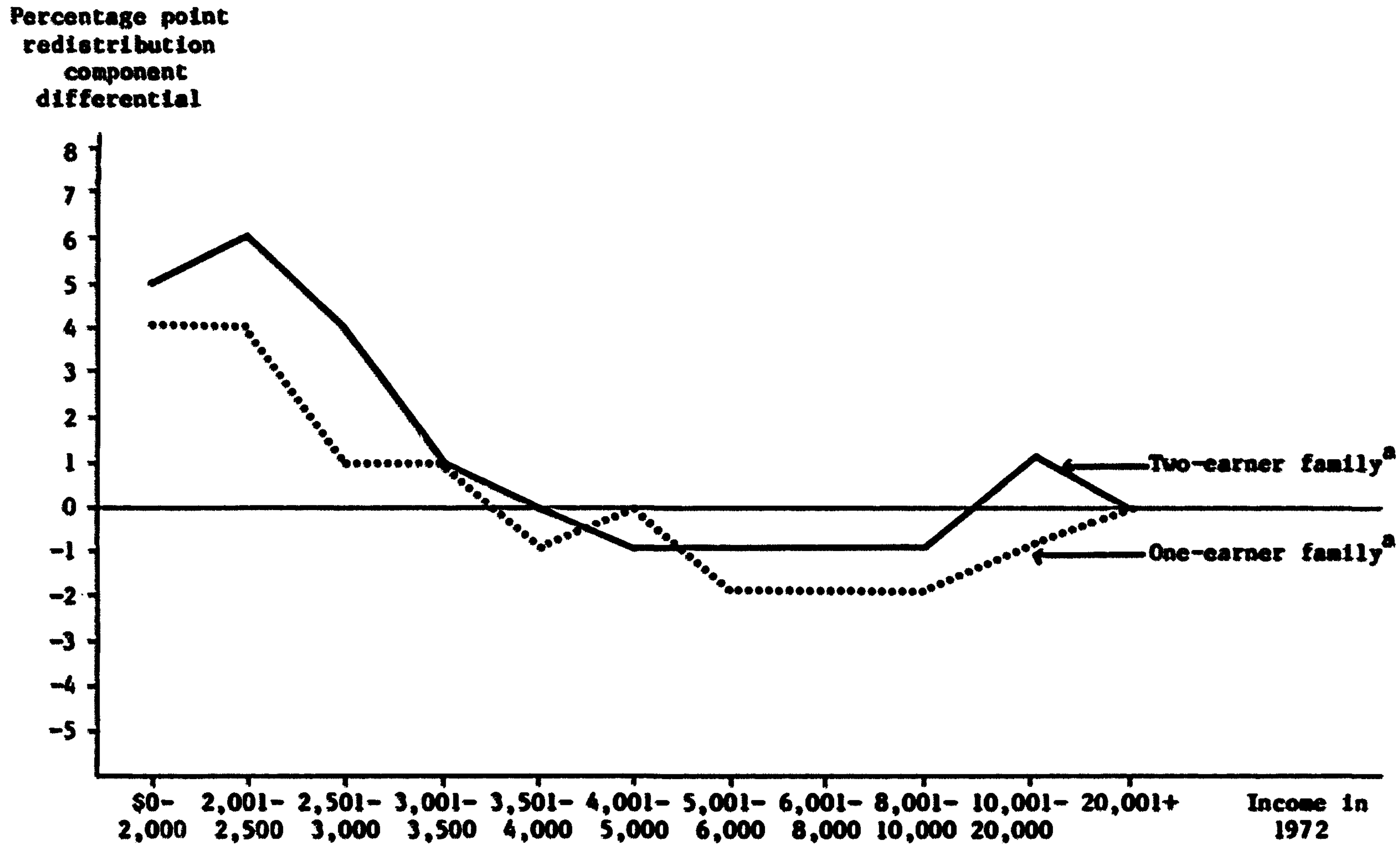

[a]Percentage of redistribution for a family income level minus percentage of redistribution for richest income level.

Figure 7.5. Progressivity of the OAI program by household type, females only




[a]Percentage of redistribution for a family income level minus percentage of redistribution for richest income level.

Figure 7.6. Progressivity of the OAI program by household type, males only




[a]Percentage of redistribution for a family income level minus percentage of redistribution for richest income level.

Figure 7.7. Progressivity of the OAI program by household type

5. The effect of social security payments on the distribution of income, both retired households only

The effect of social security benefits on the distribution of income among elderly households was examined by dividing all married couples where both members were collecting OAI benefits between 1962 and 1972 into quintile groups. Table 7.17 presents the distribution of income before and after payment of social security benefits. The distribution of personal income, exclusive of social security benefits, was highly skewed; the poorest 60 percent of the elderly population had less then 20 percent of personal income compared to the 60 percent of personal income held by the richest 20 percent of the elderly population. The addition of the husband's OAI benefits did reduce the skewedness in the distribution of income. Column 2 displays the distribution of personal income inclusive of the husband's OAI benefits, but exclusive of the wife's OAI benefits. Now, the poorest 60 percent received 30 percent of personal income, whereas the richest 20 percent received just under 50 percent of personal income. Column 3 displays the distribution of personal income after all family OAI benefits were apportioned. The distribution of personal income was, in spite of the social security program, skewed in favor of the richest quintile, but the program did increase the relative share of personal income received by the poorest 60 percent of the elderly. After receipt of all family OAI benefits, the poorest 60 percent had 34 percent of personal income compared to 45 percent of personal income received by the richest quintile. Also, the husband's share of OAI benefits had the greatest redistributional impact. This is expected since the absolute size of the male's OAI benefit generally

Table 7.17. Distribution of income for both retired population, before and after payment of social security benefits

| | (1) Distribution of personal income before social security | (2) Distribution of personal income after husband's OAI benefits | (3) Distribution of personal income after family OAI benefits | Mean personal income | | |
|---|---|---|---|---|---|---|
| | | | | Without social security | With husband benefit only | With family OAI benefits |
| Poorest quintile | 1.0% | 6.0% | 8.0% | $210 | $2,142 | $3,141 |
| Second quintile | 6.0 | 10.0 | 11.0 | 1,305 | 3,255 | 4,331 |
| Third quintile | 11.0 | 14.0 | 15.0 | 2,535 | 4,557 | 5,635 |
| Fourth quintile | 21.0 | 21.0 | 21.0 | 4,913 | 6,877 | 7,973 |
| Richest quintile | 61.0 | 49.0 | 45.0 | 14,337 | 16,270 | 17,400 |
| Total | 100.0 | 100.0 | 100.0 | | | |

exceeded the female's OAI benefit because of the male's higher average earnings and stronger labor force attachment, and because females typically collect auxiliary benefits which are 50 percent of the male's PIA.

Table 7.18 looks at the distribution of social security benefits by percentage share. Married couples in the sample received approximately $4.2 million in OAI benefits in 1972, of which 65 percent were paid to male beneficiaries and 35 percent were paid to female beneficiaries.

Overall, social security benefits were proportionally distributed to households, male beneficiaries, and female beneficiaries. Nevertheless, the roughly proportional distribution of OAI benefits significantly improved the level of personal income for the poorest 60 percent of the elderly population. The poorest quintile received 19.3 percent of all social security benefits paid to both retired, married couples in 1972, which increased its level of personal income by 1,394 percent.

In conclusion, column 1 of Table 7.17 indicates that the distribution of personal income before social security was sharply skewed in favor of the richest income quintile. The single-period analysis of OAI transfers showed that, although the distribution of personal income after the addition of social security benefits was not distributed particularly evenly, there had been a relatively small change toward increasing income equality as a result of the program's intergenerational transfer

Table 7.18. Distribution of social security benefits to both retired population by percentage share

| Quintile group | All OAI benefits | Male beneficiaries' OAI benefits | Female beneficiaries' OAI benefits | Percentage gain in personal income | | |
|---|---|---|---|---|---|---|
| | | | | Household | Male beneficiary | Female beneficiary |
| Total | 4,231,935 | 2,732,532 | 1,499,403 | | | |
| Percentage | 100.0 | 64.50 | 35.43 | | | |
| Poorest quintile | 19.3% | 19.7% | 18.6% | 1,394 | 919 | 475 |
| Second quintile | 19.9 | 19.9 | 20.0 | 232 | 149 | 83 |
| Third quintile | 20.4 | 20.6 | 20.1 | 122 | 80 | 42 |
| Fourth quintile | 20.2 | 20.1 | 20.3 | 62 | 40 | 22 |
| Richest quintile | 20.1 | 19.7 | 21.0 | 21 | 13 | 8 |
| Total | 100.0 | 100.0 | 100.0 | | | |

mechanism and income-smoothing feature.[1] A closer look at the disbursement of OAI benefits (Table 7.18) showed that benefits were, at best, proportionally distributed across quintile groups, but the largest relative gains in the level of personal income, before and after social security benefits, were realised by the poorest 60 percent of the elderly population.

The use of single-period analysis to assess the distributional impact of social security is insightful, but it can be very misleading since it fails to distinguish between the intergenerational transfer and income-smoothing features of the program. Because benefits are contingent on past OAI contributions, they are a mixture of the return on past contributions, redistribution within a retirement cohort, and redistribution across generations. The following tables in this section focus on the distributional impact of the intergenerational transfer mechanism only; that is, the income-smoothing feature has been stripped away by use of type-6 annuity counterfactuals. Table 7.19 presents the distribution of redistribution components by quintile group, controlling for family type and sex. The distribution of income before and after apportioning the redistribution component is displayed in Table 7.20.

Similar to the distribution pattern of social security benefits, the redistribution components were distributed roughly equally across

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[1]Recall, the social security program has two primary features: 1) an income-smoothing feature whereby workers transfer a fraction of their labor earnings to their retirement years by participating in the program during their earning years, and 2) an intergenerational transfer feature whereby income is transferred from the current working generation to the currently retired population.

Table 7.19. Distribution of redistribution components by quintile group controlling for family type and sex

| | Two-earner | | | | One-earner | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) Female's share of redistribution[a] | (2) Male's share of redistribution[a] | (3) Household's share of redistribution[a] | Population distribution | (4) Female's share of redistribution[a] | (5) Male's share of redistribution[a] | (6) Household's share of redistribution[a] | Population distribution |
| Poorest quintile | 16.7% | 18.1% | 17.5% | 18% | 20.8% | 21.1% | 21.0% | 22% |
| Second quintile | 20.1 | 20.5 | 20.4 | 20 | 19.7 | 19.4 | 19.4 | 20 |
| Third quintile | 20.3 | 22.0 | 21.3 | 22 | 18.9 | 19.4 | 19.3 | 19 |
| Fourth quintile | 21.0 | 20.4 | 20.6 | 21 | 19.6 | 19.6 | 19.6 | 20 |
| Richest quintile | 21.9 | 18.9 | 20.1 | 20 | 21.0 | 20.5 | 20.6 | 20 |
| Total | 674,798 | 1,075,773 | 1,767,702 | 614 | 560,920 | 1,416,564 | 1,989,561 | 780 |

[a]Redistribution component calculations are based on type-6, earnings-adjusted counterfactual.

Table 7.20. Distribution of income for married, both retired population before and after apportioning the redistribution component

| Quintile group | All households | | Two-earner household | | | One-earner household | | |
|---|---|---|---|---|---|---|---|---|
| | (1) Distribution of personal income before social security | (2) Distribution of personal income after RC[a] | (3) Distribution of personal income before RC[a] | (4) Distribution of personal income after RC[a] | (5) Net effect | (6) Distribution of personal income before RC[a] | (7) Distribution of personal income after RC[a] | (8) Net effect |
| Poorest quintile | 1.0% | 7.7% | 0.7% | 7.1% | +6.4 | 1.1% | 8.2% | +7.1 |
| Second quintile | 6.0 | 10.9 | 5.6 | 11.2 | +5.6 | 5.6 | 10.6 | +5.0 |
| Third quintile | 11.0 | 14.3 | 11.6 | 15.2 | +3.6 | 10.4 | 13.5 | +3.1 |
| Fourth quintile | 21.0 | 20.7 | 21.1 | 20.9 | -0.2 | 21.1 | 20.6 | -0.5 |
| Richest quintile | 61.0 | 46.4 | 60.0 | 45.6 | -14.4 | 61.8 | 47.1 | -14.7 |

[a]Redistribution component calculations are based on type-6, earnings-adjusted counterfactuals.

quintile groups, independent of family type and sex (Table 7.19). It is interesting to note that 22 percent of the poorest one-earner households received approximately 21 percent of all intergenerational transfers to males and females in one-earner households. Column 4 in Table 7.19 indicates that spousal benefits were, at best, proportionally distributed to dependent spouses of male workers and, therefore, were not distributed principally to needy dependent spouses as intended by the spousal benefit provision.

Table 7.20 displays the distribution for married, both retired population before and after apportioning the redistribution components. Comparing column 2 of Table 7.20 and column 3 of Table 7.17, it is clear that single-period analysis tends to overstate the true distributional impact of the OAI program. The intergenerational transfer mechanism did increase income equality but not to the extent that single-period analysis alleges or the "social adequacy" objective would seem to dictate.

## B. Regression Results

The regression results reported in this section are based on the eight models described in Chapter VI. There are four permutations of the generalized single model labeled 1, 2, 3, and 4. Recall that the specified models have identical independent variables but different dependent variables measuring the extent of redistribution. Similarly, there are four versions of the married model each having identical independent variables, but, again, different measures of redistribution were used as dependent variables.

In Chapter VI, the independent and dependent variables were defined and explained. The regression results presented in this section are organized as follows: 1) findings for the single model; 2) findings for the married model; and 3) summary of findings.

1. Single models

The expected signs of the coefficients were discussed in Chapter VI and are summarized in Table 7.21. Linear and loglinear models were estimated, in addition to the quadratic model, but the quadratic variables LTEAR2 and SERLEN2 were found to be jointly significant in all permutations of the generalized single model, although the quadratic terms, when taken separately, were not always found to be statistically significant. Summary statistics for the independent variables employed in the single model appear in Table 7.22. As might be expected, there was evidence of correlation between the labor force experience variables (LTEAR and SERLEN). The estimated correlation coefficient was 0.91 and it is statistically significant at the one percent level. In spite of the strong correlation between the two labor force variables, the estimated coefficients on LTEAR and SERLEN were significantly different from zero at a one percent level for all permutations of the single model. At present, there is no obvious solution to this multicollinearity problem without introducing a new statistical problem, specifically, a specification error. However, the construction of a larger,

Table 7.21. Single regression model variables and expected coefficient signs for models 1, 2, 3, and 4 by survivorship assumption

| Independent variable | Dependent variables in models 1, 2, 3, and 4: Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| LTEAR | Negative | Negative | Negative or positive |
| LTEAR2 | Positive | Positive | Negative or positive |
| SEX | Negative or positive | Positive | Positive |
| RACE | Negative or positive | Negative | Negative |
| SERLEN | Negative | Negative | Negative |
| SERLEN2 | Negative or positive | Negative or positive | Negative or positive |
| RAGER1 | Negative | Negative | Negative |
| RAGER2 | Negative | Negative | Negative |
| RAGER3 | Negative | Negative | Negative |
| RCOHORT1 | Positive | Positive | Positive |
| RCOHORT2 | Positive | Positive | Positive |
| EDU1 | Positive | Positive | Negative or positive |
| EDU2 | Negative or positive | Negative or positive | Negative or positive |
| EDU3 | Negative | Negative | Negative or positive |
| EDU4 | Negative | Negative | Negative or positive |

Table 7.22. Summary statistics for independent variables employed in the single regression models

| Variable | Mean | Standard deviation | Minimum | Maximum |
|---|---|---|---|---|
| LTEAR | 142,211 | 112,408 | 0 | 434,835 |
| SEX | 0.61 | 0.49 | 0 | 1 |
| RACE | 0.06 | 0.24 | 0 | 1 |
| SERLEN | 19.08 | 9.64 | 0 | 36 |
| RAGER1 | 0.50 | 0.50 | 0 | 1 |
| RAGER2 | 0.24 | 0.42 | 0 | 1 |
| RAGER3 | 0.11 | 0.31 | 0 | 1 |
| RCOHORT1 | 0.32 | 0.47 | 0 | 1 |
| RCOHORT2 | 0.31 | 0.46 | 0 | 1 |
| EDU1 | 0.21 | 0.41 | 0 | 1 |
| EDU2 | 0.10 | 0.31 | 0 | 1 |
| EDU3 | 0.29 | 0.46 | 0 | 1 |
| EDU4 | 0.20 | 0.40 | 0 | 1 |

more diverse data set is likely to minimize the collinearity present between the labor force variables in this small, relatively homogenous single data set.

a. Estimation of the model using the annuity counterfactuals for a nonindexed, no earnings test adjusted insurance program   As mentioned in Chapter VI, this permutation of the single model was estimated to isolate the partial effect of worker-specific characteristics on the percentage of redistribution in the absence of cost-of-living and earnings test adjustments. This narrow definition of the program allows for the isolation of the initial effect of the progressive benefit formula and the minimum benefit provision. The results for model 1 under different survivorship assumptions are reported in Table 7.23.

Looking first at the regression results for the model based on the gender-merged survivorship assumption (column 1 in Table 7.23), it is worth noting that all the coefficients for the independent variables have the predicted sign (for those independent variables with predicted signs). The coefficients on the quantitative variables LTEAR and SERLEN are significantly different from zero at a one percent level; however, the coefficients for the quadratic terms LTEAR2 and SERLEN2 were not significantly different from zero at a five percent level, although they were jointly significant at a one percent level. The coefficients on the control variables, RAGER3, RCOHORT1, RCOHORT2, and EDU4, were significantly different from zero at the one percent level. The coefficients on SEX, RACE, and EDU2 (variables with unpredicted coefficient signs) were not significantly different from zero.

Table 7.23. Single regression results: Model 1 under different survivorship assumptions[a,b]

| | Survivorship probability assumption | | |
|---|---|---|---|
| variable | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| LTEAR | -4.426[a] (4.02) | -4.166[a] (3.68) | -4.288[a] (3.34) |
| LTEAR2 | 0.219[d] (0.91) | 0.173 (0.70) | 0.176 (0.62) |
| SEX | -0.013 (0.03) | 3.063[a] (6.55) | 4.745[a] (8.94) |
| RACE | 0.971 (1.06) | 0.129 (0.14) | -0.111 (0.10) |
| SERLEN | -0.636[a] (5.17) | -0.626[a] (4.95) | -0.664[a] (4.63) |
| SERLEN2 | 0.006[C] (1.71) | 0.006[D] (1.59) | 0.006 (1.27) |
| RAGER1 | -0.500[d] (0.90) | -0.354 (0.62) | -0.354 (0.55) |
| RAGER2 | -0.600[d] (0.98) | -0.405 (0.65) | -0.323 (0.46) |
| RAGER3 | -2.126[a] (2.78) | -2.315[a] (2.95) | -2.834[a] (3.18) |
| RCOHORT1 | 9.240[a] (16.45) | 8.975[a] (15.54) | 9.634[a] (14.70) |
| RCOHORT2 | 6.116[a] (11.37) | 5.980[a] (10.81) | 6.495[a] (10.35) |
| EDU1 | 0.006 (0.01) | -0.092 (0.13) | -0.394 (0.51) |
| EDU2 | -0.800 (1.0) | -0.625 (0.76) | -0.482 (0.52) |
| EDU3 | -0.540[d] (0.86) | -0.619[d] (0.96) | -0.573 (0.78) |
| EDU4 | -2.754[a] (3.89) | -3.050[a] (4.18) | -2.535[A] (3.07) |
| INTERCEPT | 95.48[a] (88.19) | 93.92[a] (84.38) | 92.45[a] (73.21) |
| $R^2$ | .871 | .863 | .855 |
| N | 353 | 353 | 353 |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Column 2 in Table 7.23 presents regression results when sex and race survivorship differentials are accounted for in the annuity counterfactuals. All the coefficients, excluding those on RACE and EDU1, have the expected sign. The coefficient on SEX is positive and significantly different from zero at a one percent level. *Ceteris paribus*, women can expect a redistribution component 3.06 percentage points larger than men because of their relatively longer life expectancies, on average. Contrary to expected results, nonwhites, after accounting for their shorter life expectancies, can expect a redistribution component 0.129 percentage points larger than whites, *ceteris paribus*.

Regression results for model 1 adjusting for socioeconomic differentials in survivorship are presented in column 3 in Table 7.23. After accounting for sex, race, marital status, education, and income differentials in survivorship, the OAI program was still found to be progressive; that is, the coefficient on LTEAR is negative and significantly different from zero at a one percent level, and, although all coefficients on the education variables are negative, only EDU4 is significantly different from zero at a one percent level. Also, the coefficient on RACE is negative, but not statistically significant.

The overall effect of accounting for differential life expectancies, in most cases, is slight. Clearly, from the size of the coefficient on SEX, women receive a significantly larger redistribution component when their relatively longer life expectancy is accounted for in their actuarially fair retirement insurance payment.

b. <u>Estimation of the model using the annuity counterfactual for an indexed, no earnings test adjusted insurance program</u> The dependent variable employed in this version of the single model is the redistribution residual, in percentage terms, assuming the retiree purchased an indexed, no earnings test annuity with her accumulated OAI contributions on the date of retirement. The variation in the residual is once again explained by the quadratic model with 12 independent variables. The estimated coefficients for model 2 by survivorship assumption appear in Table 7.24.

In column 1, coefficients on LTEAR, SERLEN, and SERLEN2 have the correct sign (those with predicted signs) and are significantly different from zero at a one percent level. And the coefficient on LTEAR2 has the correct sign and is significantly different from zero at a five percent level. All the control variables have the correct sign, and coefficients on RAGER3, RCOHORT1, RCOHORT2, and EDU4 are significantly different from zero at a one percent level.

Regression results for model 2 accounting for sex and race differentials in survivorship are shown in column 2 of Table 7.24. The coefficients have the expected sign (those with predicted signs) except for RAGER1 and EDU1. Incorporating indexing and survivorship differentials by race and sex into the measure of redistribution results in coefficients on the age of retirement variables that are mixed in sign but small in size for RAGER1 and RAGER2. The positive coefficient on RAGER1 suggests that persons will maximize the percentage of redistribution by retiring at ages 62 to 64 when lifetime contributions

Table 7.24. Single regression results: Model 2, under different survivorship assumptions[a,b]

| variable | Survivorship probability assumption: Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| LTEAR | -4.871[a] | -4.523[a] | -4.884[a] |
| | (5.48) | (4.99) | (4.31) |
| LTEAR2 | 0.334[b] | 0.276[c] | 0.366[d] |
| | (1.72) | (1.39) | (1.48) |
| SEX | -0.072 | 3.101[a] | 5.192[a] |
| | (0.19) | (8.29) | (11.12) |
| RACE | 0.711 | -0.117 | -3.487[a] |
| | (0.96) | (0.16) | (3.71) |
| SERLEN | -0.662[a] | -0.630[a] | -0.676[a] |
| | (6.68) | (6.22) | (5.35) |
| SERLEN2 | 0.009[A] | 0.009[A] | 0.007[C] |
| | (3.14) | (2.90) | (1.95) |
| RAGER1 | -0.054 | 0.105 | 0.182 |
| | (0.12) | (0.23) | (0.32) |
| RAGER2 | -0.467[d] | -0.274 | -0.116 |
| | (0.95) | (0.55) | (0.19) |
| RAGER3 | -2.415[a] | -2.50[a] | -3.012[a] |
| | (3.92) | (3.97) | (3.84) |
| RCOHORT1 | 6.252[a] | 5.90[a] | 6.529[a] |
| | (13.82) | (12.76) | (11.32) |
| RCOHORT2 | 4.066[a] | 3.876[a] | 4.391[a] |
| | (9.37) | (8.75) | (7.95) |
| EDU1 | 0.087 | -0.055 | -0.303 |
| | (0.16) | (0.10) | (0.44) |
| EDU2 | -0.546 | -0.400 | -0.309 |
| | (0.85) | (0.61) | (0.38) |
| EDU3 | -0.467[d] | -0.522[d] | -0.535 |
| | (0.92 | (1.01) | (0.83) |
| EDU4 | -2.189[a] | -2.41[a] | -2.069[A] |
| | (3.83) | (4.13) | (2.84) |
| INTERCEPT | 97.55[a] | 95.92[a] | 94.395[a] |
| | (111.74) | (107.68) | (84.90) |
| $R^2$ | .881 | .870 | .853 |
| N | 353 | 353 | 353 |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

are used to purchase inflation and income insurance. This result may be more reflective of the way annuity benefits were indexed after retirement and the population distribution of the single data set than of the actual structure of the OAI program. This will be discussed further in subsection 3.

Similar results are obtained from the use of socioeconomic-adjusted survivorship probabilities, except the coefficients for SEX and RACE were found to be more statistically significant.

c. Estimation of the single model using the annuity counterfactual for a nonindexed, earnings test adjusted insurance program The nominal annuity benefit employed to calculate the dependent variable was adjusted by the OAI earning test formula for post-retirement earnings in excess of \$1,680. The quadratic model had less explanatory power, as reflected by the significantly smaller $R^2$, because 65 percent of persons with post-retirement earnings in excess of \$1,680 would have received zero annuity benefits for 1972, resulting in redistribution components equal to 100 percent.

The estimated coefficients in column 1 of Table 7.25 have the predicted sign with exception of EDU1; however, only the coefficients for LTEAR, SERLEN, RCOHORT1, and RCOHORT2 are significantly different from zero at a one percent level and the coefficient for RAGER3 is significantly different from zero at a five percent level. Nonwhites and persons with 0-7 years of education received slightly less redistribution from the OAI program relative to whites and persons with eight years of education, respectively.

Table 7.25. Single regression results: Model 3 under different survivorship assumptions[a,b]

| variable | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| | Survivorship probability assumption | | |
| LTEAR | $-6.586^{a}$ | $-6.208^{a}$ | $-6.460^{a}$ |
| | (3.55) | (3.34) | (3.18) |
| LTEAR2 | $0.516^{d}$ | $0.411^{d}$ | 0.451 |
| | (1.27) | (1.01) | (1.01) |
| SEX | 0.393 | $3.206^{a}$ | $4.789^{a}$ |
| | (0.51) | (4.18) | (5.72) |
| RACE | -0.048 | -0.852 | -1.201 |
| | (0.03) | (0.55) | (0.71) |
| SERLEN | $-0.625^{a}$ | $-0.605^{a}$ | $-0.619^{a}$ |
| | (3.02) | (2.91) | (2.73) |
| SERLEN2 | $0.012^{C}$ | $0.011^{C}$ | $0.01^{b}$ |
| | (1.89) | (1.79) | (1.46) |
| RAGER1 | -0.513 | -0.301 | -0.291 |
| | (0.55) | (0.32) | (0.28) |
| RAGER2 | -0.317 | -0.211 | -0.144 |
| | (0.31) | (0.21) | (0.13) |
| RAGER3 | $-2.839^{b}$ | $-2.918^{b}$ | $-3.431^{a}$ |
| | (2.21) | (2.26) | (2.44) |
| RCOHORT1 | $6.50^{a}$ | $6.26^{a}$ | $6.753^{a}$ |
| | (6.87) | (6.61) | (6.53) |
| RCOHORT2 | $3.109^{a}$ | $2.993^{a}$ | $3.343^{a}$ |
| | (3.44) | (3.30) | (3.37) |
| EDU1 | -0.295 | -0.482 | -0.824 |
| | (0.26) | (0.43) | (0.67) |
| EDU2 | -0.072 | -0.069 | 0.233 |
| | (0.05) | (0.05) | (0.16) |
| EDU3 | -0.785 | -0.802 | -0.736 |
| | (0.74) | (0.75) | (0.63) |
| EDU4 | $-1.783^{c}$ | $-2.133^{b}$ | $-1.689^{D}$ |
| | (1.50) | (1.78) | (1.29) |
| INTERCEPT | $97.93^{a}$ | $96.42^{a}$ | $95.11^{a}$ |
| | (53.78) | (52.81) | (47.69) |
| $R^2$ | .628 | .629 | .637 |
| N | 353 | 353 | 353 |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Again, the introduction of differentials in survivorship, be it sex-race or socioeconomic, increases the size and significance of the coefficients for RACE and SEX. With the earnings test adjustment of annuity benefits, the level of education variables follow a curious path when mortality differentials are introduced. First, including mortality differentials by sex and race in the annuity counterfactual tends to increase the negative redistributional differential between persons with less than eight or more than 11 years of education relative to persons with eight years of education. But, there is a slight narrowing of the redistributional differential between persons with 9-11 years of education relative to persons with only eight years of schooling when sex and race differentials are reflected in mortality rates. Further disaggregation of mortality rates by marital status, income, and education levels tends to strengthen the tendency of the sex and race adjustments for EDU1 only. For all other education categories, the redistributional differential is narrowed, and, for EDU2, the differential sign is positive. This suggests that the earnings test slightly weakens the program's progressivity, which is consistent with the smaller coefficients for LTEAR in columns 2 and 3 relative to column 1.

d. Estimation of the single model using the annuity counterfactual for an indexed, earnings test adjusted insurance program Similar results are obtained with this final permutation of the generalized single model, where the dependent variable is based on an annuity counterfactual promising to pay a real stream of benefits for the life of the annuitant and some or all benefits are forfeited if post-retirement earnings exceed $1,680. (The fraction forfeited depends on the size of the annuity benefit and the amount of earnings over $1,680.) The regression results are reported in Table 7.26.

With the notable exception of the coefficients for the education variables EDU1 and EDU2 in column 1, RAGER1 and EDU1 in column 2, and RAGER1 in column 3, all the coefficients have the expected sign. In column 1, the coefficients on EDU1 and EDU2 are negative and positive, respectively, indicating that persons with less than eight years of education received less, and persons with 9-11 years of education received more, redistribution per dollar of OAI benefit relative to persons with eight years of schooling. The redistributional differential generally increases with the incorporation of disaggregated mortality differentials.

e. <u>Comparison of models 1, 2, and 4 controlling for differential survivorship probabilities</u> In the previous subsections, the effect of differential mortality on the estimated coefficients across permutations of the generalized single model was examined. This subsection focuses on the effect of different program features on the size and sign of the estimated parameters, holding the survivorship assumption constant. The coefficient estimates for models 1, 2, and 4 for the gender-merged and socioeconomic-adjusted survivorship probability assumptions are reproduced in Tables 7.27 and 7.28, respectively.

Looking first at the coefficients in Table 7.27, it is interesting to note that benefit indexing and earnings test adjustments, when

Table 7.26. Single regression results: Model 4, under different survivorship assumptions[a,b]

| variable | Survivorship probability assumption | | |
|---|---|---|---|
| | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| LTEAR | -6.589[a] | -6.101[a] | -6.639[a] |
| | (4.11) | (3.87) | (3.65) |
| LTEAR2 | 0.528[c] | 0.439[d] | 0.532[D] |
| | (1.51) | (1.17) | (1.34) |
| SEX | 0.284 | 3.164[a] | 5.156[a] |
| | (0.43) | (4.87) | (6.87) |
| RACE | -0.103 | -0.943 | -4.362[a] |
| | (0.08) | (0.72) | (2.89) |
| SERLEN | -0.661[a] | -0.628[a] | -0.649[a] |
| | (3.70) | (3.57) | (3.20) |
| SERLEN2 | 0.014[A] | 0.014[A] | 0.012[B] |
| | (2.67) | (2.62) | (1.95) |
| RAGER1 | -0.129 | 0.063 | 0.149 |
| | (0.16) | (0.08) | (0.16) |
| RAGER2 | -0.283 | -0.227 | -0.074 |
| | (0.32) | (0.26) | (0.07) |
| RAGER3 | -3.051[a] | -3.044[a] | -3.576[a] |
| | (2.75) | (2.79) | (2.83) |
| RCOHORT1 | 3.851[a] | 3.598[a] | 3.989[a] |
| | (4.72) | (4.48) | (4.30) |
| RCOHORT2 | 1.435[b] | 1.331[b] | 1.598[b] |
| | (1.84) | (1.73) | (1.80) |
| EDU1 | -0.136 | -0.392 | -0.701 |
| | (0.14) | (0.41) | (0.64) |
| EDU2 | 0.115 | 0.138 | 0.293 |
| | (0.10) | (0.12) | (0.22) |
| EDU3 | -0.685 | -0.701 | -0.711 |
| | (0.75) | (0.72) | (0.68) |
| EDU4 | -1.417[c] | -1.723[b] | -1.441[a] |
| | (1.38) | (1.70) | (1.23) |
| INTERCEPT | 99.747[a] | 98.229[a] | 96.934[a] |
| | (63.42) | (63.51) | (54.26) |
| $R^2$ | .624 | .624 | .636 |
| N | 353 | 353 | 353 |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Table 7.27. Single regression results: Comparison of models 1, 2, and 4 using gender-merged survivorship probabilities

| Variable | Model 1 | Model 2 | Model 4 |
|---|---|---|---|
| LTEAR | -4.426 | -4.871 | -6.589 |
| LTEAR2 | 0.219 | 0.334 | 0.528 |
| SEX | -0.013 | -0.072 | 0.284 |
| RACE | 0.971 | 0.711 | -0.103 |
| SERLEN | -0.636 | -0.662 | -0.661 |
| SERLEN2 | 0.006 | 0.009 | 0.014 |
| RAGER1 | -0.500 | -0.054 | -0.129 |
| RAGER2 | -0.600 | -0.467 | -0.283 |
| RAGER3 | -2.126 | -2.415 | -3.051 |
| RCOHORT1 | 9.240 | 6.252 | 3.851 |
| RCOHORT2 | 6.116 | 4.066 | 1.435 |
| EDU1 | 0.006 | 0.087 | -0.136 |
| EDU2 | -0.800 | -0.546 | 0.115 |
| EDU3 | -0.540 | -0.467 | -0.685 |
| EDU4 | -2.754 | -2.189 | -1.417 |
| Intercept | 95.48 | 97.55 | 99.747 |
| $R^2$ | .871 | .881 | .628 |

Table 7.28. Single regression results: Comparison of models 1, 2, and 4 using socioeconomic-adjusted survivorship probabilities

| Variable | Model 1 | Model 2 | Model 4 |
|---|---|---|---|
| LTEAR | -4.288 | -4.884 | -6.639 |
| LTEAR2 | 0.176 | 0.366 | 0.532 |
| SEX | 4.745 | 5.192 | 5.156 |
| RACE | -0.111 | -3.487 | -4.362 |
| SERLEN | -0.664 | -0.676 | -0.649 |
| SERLEN2 | 0.006 | 0.007 | 0.012 |
| RAGER1 | -0.354 | 0.182 | 0.149 |
| RAGER2 | -0.323 | -0.116 | -0.074 |
| RAGER3 | -2.834 | -3.012 | -3.576 |
| RCOHORT1 | 9.634 | 6.529 | 3.989 |
| RCOHORT2 | 6.495 | 4.391 | 1.598 |
| EDU1 | -0.394 | -0.303 | -0.701 |
| EDU2 | -0.482 | -0.309 | 0.293 |
| EDU3 | -0.573 | -0.535 | -0.711 |
| EDU4 | -2.535 | -2.069 | -1.441 |
| Intercept | 92.45 | 94.395 | 96.934 |
| $R^2$ | .855 | .853 | .636 |

accounted for in the annuity counterfactual, do have an effect on the relationship between the independent and dependent variables as reflected in the estimated coefficients. For instance, the coefficient on the lifetime earnings measure increases in absolute size with the introduction of indexing and earning test adjustments into the annuity counterfactual. At first blush, this evidence would tend to suggest that the program becomes more progressive as the annuity counterfactual more closely approximates the OAI program. However, this generalization may be too strong in light of the observed pattern on the coefficients for LTEAR2 and the education variables. The coefficient for LTEAR2 enters with a positive sign in column 1 and increases across the model, offsetting the strength of the negative coefficient on LTEAR. Likewise, the coefficients on the education variable show a weakening of progressivity across the models. The coefficient estimates for EDU1 across the models show a withering away of the redistributional gains for persons with 0-7 years of education relative to persons with eight years of schooling. The redistributional losses associated with education levels of 13 or more years of education are reduced, and for education levels 9-11 the loss not only diminishes but becomes a gain when the earning test is added to the annuity counterfactual.

A few additional patterns across models are worth mentioning. The sign change on the estimated coefficient for SEX with the accounting for the earnings test suggests that women were more likely to continue working after retirement and, as a result, women tended to have slightly larger redistribution components. The pattern on the coefficient for

RACE, on the other hand, suggests that the redistributional gains of nonwhites are reduced under indexing and, with the addition of an earnings test, nonwhites receive slightly less redistribution when compared to their white counterparts. The last, and perhaps the most dramatic, pattern to be mentioned concerns the estimated coefficients on the retirement cohort variables, RCOHORT1 and RCOHORT2. The redistribution gains for persons retiring in 1962-1965 and 1966-1968 relative to the 1969-1972 retirement cohort consistently diminish across models.

Similar results are observed using socioeconomic-adjusted probabilities (see Table 7.28). It is interesting to note that females received slightly more redistribution from an indexed system relative to males, again, because of their longer life expectancies. Alternatively, nonwhites are made significantly worse off, in terms of the reduced share of redistribution from an indexed system, relative to whites because of race differentials in survivorship (compare columns 2 and 3).

2. Married models

Reported regression results are based on the estimation of four permutations of the generalized married quadratic model. Linear and loglinear models were estimated, but the quadratic variables FLTEAR2, SERLEN2, and _SERLEN2 were found to be jointly, although only FLTEAR2 was found to be separately, significant in all permutations of the generalized model. The expected signs for all 24 independent variables are summarized in Table 7.29, and summary statistics for each independent

Table 7.29. Married regression model variables and expected coefficient signs for models 5, 6, 7, and 8 by survivorship assumption

| Independent variable | Dependent variables in models 5, 6, 7, and 8 | | |
|---|---|---|---|
| | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| FLTEAR | Negative | Negative | Negative or positive |
| FLTEAR2 | Positive | Positive | Negative or positive |
| RACE | Negative or positive | Negative | Negative |
| SERLEN | Negative | Negative | Negative |
| _SERLEN | Negative | Negative | Negative |
| SERLEN2 | Negative or positive | Negative or positive | Negative or positive |
| _SERLEN2 | Negative or positive | Negative or positive | Negative or positive |
| RAGER1 | Negative | Negative | Negative |
| RAGER2 | Negative | Negative | Negative |
| RAGER3 | Negative or positive | Negative or positive | Negative or positive |
| _RAGER1 | Negative | Negative | Negative |
| _RAGER2 | Negative | Negative | Negative |
| RCOHORT1 | Positive | Positive | Positive |
| RCOHORT2 | Positive | Positive | Positive |
| _RCOHORT1 | Positive | Positive | Positive |
| _RCOHORT2 | Positive | Positive | Positive |
| EDU1 | Positive | Positive | Negative or positive |
| EDU2 | Negative or positive | Negative or positive | Negative or positive |
| EDU3 | Negative | Negative | Negative or positive |
| EDU4 | Negative | Negative | Negative or positive |
| _EDU1 | Positive | Positive | Negative or positive |
| _EDU2 | Negative or positive | Negative or positive | Negative or positive |
| _EDU3 | Negative | Negative | Negative or positive |
| _EDU4 | Negative | Negative | Negative or positive |

variable appear in Table 7.30. There was evidence of correlation between the service length variables within a household, but collinearity was not a problem between the lifetime earnings measure (FLTEAR) and service length variables (SERLEN, _SERLEN). The correlation coefficient on the service length variables SERLEN and _SERLEN was relatively small, 0.33, but significantly different from zero at the five percent level.

a. Estimation of the model using the annuity counterfactual for a nonindexed, no earnings test adjusted insurance program  As discussed in Chapter VI, the annuity counterfactual used to determine the percentage of redistribution was based on the assumption that the retirement candidate purchased a life annuity that promised payment of a nominal stream of income for life and the size of the benefit payment was invariant to post-retirement earnings. Then, the quadratic model with 24 independent variables was estimated to isolate the partial effect of household-specific characteristics on the percentage of redistribution for the household. The results for model 5 under different survivorship assumptions are presented in Table 7.31.

In the regression for the gender-merged survivorship probabilities (column 1), all independent variables have the predicted sign, with the exception of SERLEN, RAGER3, EDU1, EDU4, _EDU1, _EDU3, and _EDU4. Of those variables with the predicted sign, only FLTEAR, FLTEAR2, _RAGER1, RCOHORT1, RCOHORT2, _RCOHORT1, and _RCOHORT2 have estimated coefficients that are significantly different from zero at a five percent level. And, of those variables with the unpredicted sign, only the coefficient for RAGER3 is significantly different from zero at a one percent level.

Table 7.30. Summary statistics for independent variables employed in the married regression models

| Variable | Mean | Standard deviation | Minimum | Maximum |
|---|---|---|---|---|
| FLTEAR | 241,996 | 155,621 | 0 | 805,200 |
| RACE | 0.02 | 0.13 | 0 | 1.00 |
| SERLEN | 6.43 | 8.41 | 0 | 35.00 |
| _SERLEN | 21.50 | 10.14 | 0 | 36.00 |
| RAGER1 | 0.76 | 0.43 | 0 | 1 |
| RAGER2 | 0.11 | 0.31 | 0 | 1 |
| RAGER3 | 0.10 | 0.31 | 0 | 1 |
| _RAGER1 | 0.42 | 0.49 | 0 | 1 |
| _RAGER2 | 0.26 | 0.44 | 0 | 1 |
| RCOHORT1 | 0.26 | 0.44 | 0 | 1 |
| RCOHORT2 | 0.29 | 0.45 | 0 | 1 |
| _RCOHORT1 | 0.36 | 0.48 | 0 | 1 |
| _RCOHORT2 | 0.31 | 0.46 | 0 | 1 |
| EDU1 | 0.18 | 0.38 | 0 | 1 |
| EDU2 | 0.17 | 0.38 | 0 | 1 |
| EDU3 | 0.25 | 0.44 | 0 | 1 |
| EDU4 | 0.16 | 0.36 | 0 | 1 |
| _EDU1 | 0.23 | 0.42 | 0 | 1 |
| _EDU2 | 0.16 | 0.37 | 0 | 1 |
| _EDU3 | 0.17 | 0.38 | 0 | 1 |
| _EDU4 | 0.15 | 0.36 | 0 | 1 |

Table 7.31. Married regression results: Model 5 under different survivorship assumptions[a,b]

| Variable | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| | Survivorship probability assumption | | |
| FLTEAR | -4.071[a] | -3.994[a] | -4.002[A] |
| | (26.81) | (26.72) | (26.73) |
| FLTEAR2 | 0.154[a] | 0.152[a] | 0.157[A] |
| | (5.80) | (5.82) | (6.00) |
| RACE | -0.305 | -0.718[d] | 0.977[a] |
| | (0.53) | (1.27) | (1.73) |
| SERLEN | 0.007 | 0.008 | 0.008 |
| | (0.25) | (0.29) | (0.29) |
| _SERLEN | -0.043[c] | -0.042[c] | -0.042[c] |
| | (1.49) | (1.51) | (1.49) |
| SERLEN2 | 0.000 | 0.000 | 0.000 |
| | (0.01) | (0.01) | (0.07) |
| _SERLEN2 | 0.001 | 0.001 | 0.001 |
| | (0.95) | (0.96) | (1.01) |
| RAGER1 | -0.056 | 0.038 | -0.044 |
| | (0.25) | (0.17) | (0.20) |
| RAGER2 | -0.429[c] | -0.460[c] | -0.528[b] |
| | (1.43) | (1.55) | (1.78) |
| RAGER3 | 0.774[a] | 0.779[a] | 0.849[a] |
| | (2.93) | (3.05) | (3.32) |
| _RAGER1 | -0.534[a] | -0.528[a] | -0.546[a] |
| | (3.05) | (3.37) | (3.16) |
| _RAGER2 | -0.291[c] | -0.273[c] | -0.295[c] |
| | (1.50) | (1.43) | (1.54) |
| RCOHORT1 | 1.874[a] | 1.746[a] | 1.788[a] |
| | (8.39) | (7.94) | (8.12) |
| RCOHORT2 | 1.493[a] | 1.421[a] | 1.444[a] |
| | (7.51) | (7.26) | (7.37) |
| _RCOHORT1 | 5.404[a] | 5.394[a] | 5.359[a] |
| | (24.67) | (25.01) | (24.82) |
| _RCOHORT2 | 3.556[a] | 3.544[a] | 3.536[a] |
| | (17.65) | (17.87) | (17.80) |
| EDU1 | -0.126 | -0.118 | -0.301 |
| | (0.52) | (0.49) | (1.26) |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Table 7.31. continued

| Variable | Survivorship probability assumption | | |
|---|---|---|---|
| | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| EDU2 | -0.201 | -0.180 | 0.084 |
| | (0.84) | (0.77) | (0.36) |
| EDU3 | -0.088 | -0.070 | 0.091 |
| | (0.38) | (0.31) | (0.40) |
| EDU4 | 0.203 | 0.223 | 0.969[A] |
| | (0.74) | (0.82) | (3.58) |
| _EDU1 | -0.107 | -0.095 | -0.038 |
| | (0.49) | (0.44) | (0.18) |
| _EDU2 | 0.332[D] | 0.317[D] | 0.309[D] |
| | (1.42) | (1.37) | (1.34) |
| _EDU3 | 0.047 | 0.044 | 0.055 |
| | (0.19) | (0.18) | (0.23) |
| _EDU4 | 0.051 | 0.043 | 0.055 |
| | (0.19) | (0.16) | (0.20) |
| Intercept | 92.560[a] | 92.655[a] | 92.520[a] |
| | (208.47) | (211.94) | (211.35) |
| $R^2$ | .849 | .848 | .846 |
| N | 1,394 | 1,394 | 1,394 |

Of the six quantitative variables, only FLTEAR and FLTEAR2 explain a significant amount of the variation of the percentage of redistribution around its mean. As expected, the estimated coefficients on FLTEAR and FLTEAR2 are negative and positive, respectively, but, when taken together, there exists a negative association between the family measure of the percentage of redistribution and family lifetime earnings. The estimated coefficients for the education variables for the husband and wife are small, and they were found to be statistically insignificant, separately and jointly. However, the signs on the education variable coefficients, especially on EDU1, EDU4, _EDU1, and _EDU4, challenge the progressivity conclusion based exclusively on the overall sign of the coefficient on the family lifetime earnings measures.

The interpretation of the other independent variables is straightforward and consistent with earlier discussions for the single models, with the exception of RAGER3. The coefficient for RAGER3 is positive and it is statistically significant. This suggests that households where the woman retired after age 71 received a redistribution component that was .774 percentage points larger than households where the woman retired at age 65, <u>ceteris paribus</u>.

Next, looking at regression results in column 2, there are but minor changes in the estimated coefficients after accounting for mortality differentials by sex and race. The coefficient for RACE, while small and statistically insignificant, indicates that nonwhite households received slightly less redistribution relative to white households. The coefficient for RACE is, however, only slightly larger after adjustments are made for race differentials in survivorship. Perhaps, though, the most curious finding is the sign switching on the coefficient for RAGER1 after introducing sex and race differentials in survivorship. Now, households where women retired before age 65 and after age 72 received slightly larger redistribution components relative to households where the woman retired at age 65.

The regression results for model 5 after accounting for socio-economic differentials in survivorship are presented in column 3 of Table 7.31. The coefficients for the following variables are significantly different from zero at a five percent level: FTEAR, FLTEAR2, RACE, RAGER2, RAGER3, _RAGER1, RCOHORT1, RCOHORT2, _RCOHORT1, _RCOHORT2, and EDU4. Two interesting results should be noted. The coefficient for RACE is positive and significantly different from zero at a one percent level after controlling for race, sex, marital status, education, and income differentials in survivorship. Also, the coefficient for EDU4 is positive and significantly different from zero at a one percent level. That is, households where the woman has some college education received a redistribution component that was approximately .97 percentage points larger than households where the woman had eight years of education.

Comparisons of the results across survivorship assumptions suggest that for married households aggregate results do not significantly change, except for RACE, RAGER1, and EDU4, with mortality rate disaggregation.

b. *Estimation of the model using the annuity counterfactual for an indexed, no earnings test adjusted insurance program.* The dependent variable employed in this version of the married model is based on an annuity counterfactual promising a real stream of benefits for the life of the annuitants. The variation in the dependent variable is once again explained by the quadratic model with 24 independent variables. The estimated coefficients by survivorship assumption appear in Table 7.32.

Regression results for model 6 based on gender-merged survivorship probabilities are reported in column 1. The coefficients for FLTEAR, FLTEAR2, RAGER2, RAGER3, _RAGER1, _RAGER2, RCOHORT1, RCOHORT2, _RCOHORT1, and _RCOHORT2 are significantly different from zero at a five percent level, and they enter with the predicted sign. When the annuity promises to pay a fixed real benefit level for the life of the annuitants, the household received slightly more redistribution if the woman elected to retire prior to age 65, as reflected by the coefficient for RAGER1. The redistribution gains are larger yet for the household when the woman retired after age 71, everything else equal.

The results for the education dummy variables are mixed with all eight coefficients small. According to the signs of the coefficients for EDU1, EDU2, EDU3, and EDU4, households received slightly less redistribution when the female member had less than eight or 9-12 years of education, whereas households received slightly more redistribution when the female member had some college education relative to households where the female member had eight years of schooling. Turning to the comparable coefficients for the male member, households where the male member had nine or more years of schooling received slightly larger redistribution components (although the marginal gain decreased with extra years of schooling), whereas the opposite was true for households

Table 7.32. Married regression results: Model 6 under different survivorship assumptions[a,b]

| | Survivorship probability assumption | | |
|---|---|---|---|
| Variable | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| FLTEAR | -3.894[a] | -3.793[a] | -3.791[A] |
| | (31.03) | (30.79) | (30.90) |
| FLTEAR2 | 0.190[a] | 0.185[a] | 0.190[A] |
| | (8.62) | (8.56) | (8.84) |
| RACE | -0.197 | -0.606[c] | 1.097[a] |
| | (0.42) | (1.30) | (2.37) |
| SERLEN | 0.006 | 0.007 | 0.007 |
| | (0.27) | (0.32) | (0.32) |
| _SERLEN | -0.031[c] | -0.031[c] | -0.030[c] |
| | (1.30) | (1.32) | (1.28) |
| SERLEN2 | 0.000 | 0.000 | 0.000 |
| | (0.01) | (0.04) | (0.10) |
| _SERLEN2 | 0.000 | 0.000 | 0.001 |
| | (0.82) | (0.83) | (0.87) |
| RAGER1 | 0.081 | 0.202[d] | 0.117 |
| | (0.44) | (1.12) | (0.65) |
| RAGER2 | -0.473[b] | -0.518[b] | -0.565[a] |
| | (1.90) | (2.12) | (2.32) |
| RAGER3 | 1.113[a] | 1.119[a] | 1.167[a] |
| | (5.19) | (5.32) | (5.57) |
| _RAGER1 | -0.394[a] | -0.473[a] | -0.435[a] |
| | (2.72) | (3.32) | (3.08) |
| _RAGER2 | -0.306[b] | -0.277[b] | -0.297[b] |
| | (1.90) | (1.75) | (1.88) |
| RCOHORT1 | 1.056[a] | 0.877[a] | 0.929[a] |
| | (5.72) | (4.84) | (5.15) |
| RCOHORT2 | 0.965[a] | 0.868[a] | 0.891[a] |
| | (5.87) | (5.38) | (5.55) |
| _RCOHORT1 | 3.691[a] | 3.764[a] | 3.711[a] |
| | (20.39) | (21.18) | (20.97) |
| _RCOHORT2 | 2.345[a] | 2.379[a] | 2.357[a] |
| | (14.08) | (14.56) | (14.48) |
| EDU1 | -0.084 | -0.074 | -0.270[D] |
| | (0.42) | (0.38) | (1.37) |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Table 7.32. continued

| Variable | Survivorship probability assumption: Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| EDU2 | -0.148 | -0.123 | 0.151 |
| | (0.75) | (0.64) | (0.78) |
| EDU3 | -0.045 | -0.023 | 0.159 |
| | (0.23) | (0.16) | (0.85) |
| EDU4 | 0.140 | 0.168 | 1.054[A] |
| | (0.62) | (0.75) | (4.75) |
| _EDU1 | -0.138 | -0.122 | -0.071 |
| | (0.76) | (0.69) | (0.40) |
| _EDU2 | 0.266[D] | 0.249[D] | 0.239[D] |
| | (1.37) | (1.31) | (1.26) |
| _EDU3 | 0.058 | 0.054 | 0.064 |
| | (0.29) | (0.27) | (0.33) |
| _EDU4 | 0.020 | 0.006 | 0.009 |
| | (0.09) | (0.03) | (0.04) |
| Intercept | 94.233[a] | 94.302[a] | 94.167[a] |
| | (256.81) | (261.83) | (262.50) |
| $R^2$ | .848 | .846 | .845 |
| N | 1,394 | 1,394 | 1,394 |

where the male member had less than eight years of education when compared to households where the male member had eight years of schooling, *ceteris paribus*.

Introducing disaggregated survivorship probabilities does change some of the basic findings under the gender-merged assumption. First, looking at the sex-race disaggregated assumption in column 2 of Table 7.32, the changes are relatively minor and confined to race and sex-distinct dummy variables. The race coefficient is slightly more negative, as are the coefficients for RACER2 and _RACER1. Alternatively, the redistributional gains to households where the female retired prior to age 65 were slightly increased; however, the redistributional gains to households where the female member retired prior to 1969 were slightly reduced.

When survivorship probabilities are further disaggregated by marital status, education, and income, the coefficient estimates affected are for the variables RACE, EDU1, EDU2, EDU3, EDU4, _EDU1, and _EDU4. Clearly, the most dramatic change pertains to the coefficient for RACE; the coefficient for RACE in column 3 is positive and significantly different from zero at a one percent level. Therefore, nonwhite households received redistribution components 1.097 percentage points larger than white households, *ceteris paribus*.

Similarly speaking, the accounting for education differentials in survivorship, in addition to sex differentials, affects the estimated coefficients for EDU1, EDU2, EDU3, EDU4, _EDU1, and _EDU4. The household measure of redistribution was smaller if the female member had less than eight years of education, but it was larger if the female member had more than eight years of education. The coefficients on EDU1, EDU2, and EDU3 were small; however, the coefficient for EDU4 was positive and significantly different from zero at the one percent level. The size of the redistributional loss for households where the male member had less than eight years of schooling decreased when sex and education differentials in survivorship were introduced. However, the estimated coefficients for _EDU2 and _EDU3 were remarkably stable under different survivorship assumptions.

c. Estimation of the married model using the annuity counterfactual for the nonindexed, earnings test adjusted insurance program The dependent variable was constructed using the nominal annuity benefit counterfactual adjusted by the OAI earnings test formula. The explanatory power of the generalized married model, as reflected by the smaller $R^2$, is significantly weakened by the larger deviations in the redistribution measure for observations affected by the earnings test. Approximately ten percent of the married households were affected by the earnings test.

All the estimated coefficients in column 1 of Table 7.33 have the predicted sign, with the exception of SERLEN, SERLEN2, EDU1, EDU4, _EDU1, _EDU3, and EDU4. Of the coefficients with the correct sign, the estimates for FLTEAR, FLTEAR2, RAGER3, _RAGER2, RCOHORT1, RCOHORT2, _RCOHORT1, and _RCOHORT2 are significantly different from zero at the five percent level. Only one of the coefficients with the wrong sign is statistically significant, EDU4. The coefficients on the service length

Table 7.33. Married regression results: Model 7 under different survivorship assumptions[a,b]

| Variable | Survivorship probability assumption | | |
|---|---|---|---|
| | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| FLTEAR | -3.949[a] | -3.874[a] | -3.884[A] |
| | (16.18) | (16.16) | (16.28) |
| FLTEAR2 | 0.176[a] | 0.176[a] | 0.179[A] |
| | (4.11) | (4.13) | (4.28) |
| RACE | -0.358 | -0.770[d] | 0.920[d] |
| | (0.39) | (0.85) | (1.02) |
| SERLEN | 0.007 | 0.008 | 0.007 |
| | (0.16) | (0.18) | (0.17) |
| _SERLEN | -0.023 | -0.023 | -0.022 |
| | (0.51) | (0.50) | (0.49) |
| SERLEN2 | -0.001 | -0.001 | -0.001 |
| | (0.69) | (0.70) | (0.72) |
| _SERLEN2 | 0.001 | 0.001 | 0.001 |
| | (0.61) | (0.60) | (0.62) |
| RAGER1 | -0.224 | -0.129 | -0.201 |
| | (0.63) | (0.37) | (0.58) |
| RAGER2 | -0.122 | -0.149 | -0.207 |
| | (0.25) | (0.31) | (0.44) |
| RAGER3 | 3.40 [a] | 3.362[a] | 3.405[a] |
| | (8.15) | (8.21) | (8.35) |
| _RAGER1 | -0.354[d] | -0.399[c] | -0.362[c] |
| | (1.26) | (1.44) | (1.31) |
| _RAGER2 | -0.601[b] | -0.578[b] | -0.594[b] |
| | (1.92) | (1.88) | (1.94) |
| RCOHORT1 | 1.748[a] | 1.631[a] | 1.680[a] |
| | (4.87) | (4.62) | (4.78) |
| RCOHORT2 | 1.373[a] | 1.308[a] | 1.337[a] |
| | (4.30) | (4.17) | (4.28) |
| _RCOHORT1 | 3.794[a] | 3.803[a] | 3.784[a] |
| | (10.77) | (11.00) | (10.99) |
| _RCOHORT2 | 1.980[a] | 1.989[a] | 1.998[a] |
| | (6.11) | (6.25) | (6.31) |
| EDU1 | -0.236 | -0.226 | -0.410 |
| | (0.60) | (0.59) | (1.07) |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Table 7.33. continued

| Variable | Survivorship probability assumption: Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
|---|---|---|---|
| EDU2 | -0.319 | -0.297 | -0.043 |
| | (0.83) | (0.79) | (0.11) |
| EDU3 | -0.396[d] | -0.370[d] | -0.216 |
| | (1.07) | (1.02) | (0.60) |
| EDU4 | 0.238 | 0.249 | 0.927[B] |
| | (0.54) | (0.57) | (2.15) |
| _EDU1 | -0.680[b] | -0.657[b] | -0.602[C] |
| | (1.92) | (1.89) | (1.74) |
| _EDU2 | 0.453[D] | 0.435[D] | 0.429[D] |
| | (1.20) | (1.18) | (1.17) |
| _EDU3 | 0.000 | -0.003 | 0.020 |
| | (0.0) | (0.01) | (0.05) |
| _EDU4 | 0.616[c] | 0.603[c] | 0.598[C] |
| | (1.41) | (1.41) | (1.40) |
| Intercept | 93.934[a] | 94.001[a] | 93.856[a] |
| | (131.60) | (134.12) | (134.51) |
| $R^2$ | .619 | .618 | .619 |
| N | 1,394 | 1,394 | 1,394 |

variables (SERLEN, _SERLEN, SERLEN2, and _SERLEN2) have mixed signs and they are statistically insignificant, separately and jointly.

The introduction of disaggragated survivorship probabilities, either by sex and race or sex, race, marital status, income, and education, does not significantly affect the aggregate results, with the notable exception of RACE and the education variables.

d. Estimation of the married model using the annuity counterfactual for an indexed, earnings test adjusted insurance program The final permutation of the generalized married model was estimated to explain the variation in the redistribution component calculated using an indexed annuity counterfactual adjusted by the OAI earnings test formula. The regression results are reported in Table 7.34 by survivorship assumption.

Based on the gender-merged assumption, the estimated coefficients for FLTEAR, FLTEAR2, RAGER3, _RAGER2, RCOHORT1, RCOHORT2, _RCOHORT1, and _RCOHORT2 have the predicted sign and were significantly different from zero at a five percent level (see column 1). The coefficient for _EDU1 was significantly different from zero at a five percent level, but it did not have the predicted sign. Again, the coefficients for the education variables were mixed and statistically insignificant, separately (with the exception of _EDU1), but not jointly.

Disaggregating survivorship probabilities by race and sex resulted in only modest changes in the coefficient estimates for RACE and RAGER1 (see column 2). Further disaggregation of survivorship probabilities by marital status, income, and education, also, resulted in only modest

Table 7.34. Married regression results: Model 8 under different survivorship assumptions[a,b]

| | Survivorship probability assumption | | |
|---|---|---|---|
| Variable | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| FLTEAR | -3.786[a] | -3.688[a] | -3.689[A] |
| | (18.22) | (18.17) | (18.36) |
| FLTEAR2 | 0.205[a] | 0.200[a] | 0.206[A] |
| | (5.64) | (5.64) | (5.85) |
| RACE | -0.234 | -0.639 | 1.059[c] |
| | (0.30) | (0.83) | (1.40) |
| SERLEN | 0.005 | 0.006 | 0.005 |
| | (0.14) | (0.16) | (0.15) |
| _SERLEN | -0.012 | -0.012 | -0.010 |
| | (0.32) | (0.31) | (0.28) |
| SERLEN2 | -0.001 | -0.001 | -0.001 |
| | (0.68) | (0.68) | (0.70) |
| _SERLEN2 | 0.000 | 0.000 | 0.000 |
| | (0.45) | (0.44) | (0.44) |
| RAGER1 | -0.039 | 0.081 | 0.007 |
| | (0.13) | (0.27) | (0.02) |
| RAGER2 | -0.221 | -0.261 | -0.299 |
| | (0.54) | (0.65) | (0.75) |
| RAGER3 | 3.262[a] | 3.226[a] | 3.240[a] |
| | (9.19) | (9.30) | (9.43) |
| _RAGER1 | -0.234[d] | -0.308[c] | -0.270[d] |
| | (0.97) | (1.32) | (1.16) |
| _RAGER2 | -0.525[b] | -0.492[b] | -0.505[b] |
| | (1.97) | (1.89) | (1.96) |
| RCOHORT1 | 0.949[a] | 0.785[a] | 0.844[a] |
| | (3.10) | (2.63) | (2.85) |
| RCOHORT2 | 0.876[a] | 0.788[a] | 0.817[a] |
| | (3.22) | (2.97) | (3.11) |
| _RCOHORT1 | 2.344[a] | 2.433[a] | 2.401[a] |
| | (7.82) | (8.31) | (8.28) |
| _RCOHORT2 | 1.046[a] | 1.098[a] | 1.098[a] |
| | (3.79) | (4.08) | (4.12) |
| EDU1 | -0.171 | -0.160 | -0.355 |
| | (0.51) | (0.49) | (1.10) |

[a]t-ratios in parentheses.

[b]Significance levels (uppercase for 2-tail tests, lowercase for 1-tail tests): A, a-1%; B, b-5%, C, c-10%, D, d-20%.

Table 7.34. continued

| Variable | Survivorship probability assumption | | |
|---|---|---|---|
| | Gender-merged | Sex-race-distinct | Socioeconomic-adjusted |
| EDU2 | -0.240 | -0.215 | 0.050 |
| | (0.74) | (0.67) | (0.16) |
| EDU3 | -0.268[d] | -0.241 | -0.066 |
| | (0.85) | (0.78) | (0.21) |
| EDU4 | 0.170 | 0.188 | 0.998[A] |
| | (0.45) | (0.51) | (2.75) |
| _EDU1 | -0.606[b] | -0.579[b] | -0.529[C] |
| | (2.02) | (1.97) | (1.82) |
| _EDU2 | 0.375 | 0.352 | 0.345 |
| | (1.17) | (1.13) | (1.11) |
| _EDU3 | 0.040 | 0.036 | 0.060 |
| | (0.12) | (0.11) | (0.19) |
| _EDU4 | 0.506[c] | 0.485[c] | 0.469[D] |
| | (1.36) | (1.34) | (1.31) |
| Intercept | 95.34[a] | 95.379[a] | 95.230[a] |
| | (156.93) | (160.70) | (162.03) |
| $R^2$ | .611 | .610 | .612 |
| N | 1,394 | 1,394 | 1,394 |

changes in the parameter estimates. The estimates for model 8 employing socioeconomic-adjusted survivorship probabilities are presented in column 3 of Table 7.34. The coefficient for RACE does not have the predicted sign and is significantly different from zero at a ten percent level. The coefficient for EDU1 is generally more negative and _EDU1 less negative as survivorship probabilities are more disaggregated. The coefficient for EDU2 turns positive when mortality differentials by marital status, income, and education are included, and, more importantly, the coefficient for EDU4 is positive and significantly different from zero at a one percent level.

e. Comparison of models 5, 6, and 8 controlling for differential survivorship probabilities In this subsection, the effect of different program features on the size and sign of the estimated coefficients will be investigated, under the same survivorship assumption. In Table 7.35, the coefficients for models 5, 6, and 8 using gender-merged survivorship probabilities are presented. Comparisons of models 5, 6, and 8 findings based on socioeconomic-adjusted survivorship probabilities appear in Table 7.36.

Most of the coefficient estimates are remarkably stable across program features, but some important trends are observed. First, the combined effect of FLYEAR and FLYEAR2 shows a weakening of the program's progressivity when the annuity counterfactual includes indexing and the earnings test. Second, the coefficient for RAGER1 is positive when benefit indexing is included in the annuity counterfactual, but becomes negative when, in addition to indexing, the earnings test is adopted.

Table 7.35. Married regression results: Comparison of models 5, 6, and 8 using gender-merged survivorship probabilities

| Variable | (1) Model 5 | (2) Model 6 | (3) Model 8 |
|---|---|---|---|
| FLTEAR | -4.071 | -3.894 | -3.786 |
| FLTEAR2 | 0.154 | 0.190 | 0.205 |
| RACE | -0.305 | -0.197 | -0.234 |
| SERLEN | 0.007 | 0.006 | 0.005 |
| _SERLEN | -0.043 | -0.031 | -0.012 |
| SERLEN2 | 0.000 | 0.000 | -0.001 |
| _SERLEN2 | 0.001 | 0.000 | 0.000 |
| RAGER1 | -0.056 | 0.081 | -0.039 |
| RAGER2 | -0.429 | -0.473 | -0.221 |
| RAGER3 | 0.774 | 1.113 | 3.262 |
| _RAGER1 | -0.534 | -0.394 | -0.234 |
| _RAGER2 | -0.291 | -0.306 | -0.525 |
| RCOHORT1 | 1.874 | 1.056 | 0.949 |
| RCOHORT2 | 1.493 | 0.965 | 0.876 |
| _RCOHORT1 | 5.404 | 3.691 | 2.344 |
| _RCOHORT2 | 3.556 | 2.345 | 1.046 |
| EDU1 | -0.126 | -0.084 | -0.171 |
| EDU2 | -0.201 | -0.148 | -0.240 |
| EDU3 | -0.088 | -0.045 | -0.268 |
| EDU4 | 0.203 | 0.140 | 0.170 |
| _EDU1 | -0.107 | -0.138 | -0.606 |
| _EDU2 | 0.332 | 0.266 | 0.375 |
| _EDU3 | 0.047 | 0.058 | 0.040 |
| _EDU4 | 0.051 | 0.020 | 0.506 |
| Intercept | 92.560 | 94.233 | 95.34 |
| $R^2$ | 0.849 | 0.848 | 0.611 |

Table 7.36. Married regression results: Comparison of models 5, 6, and 8 using socioeconomic-adjusted survivorship probabilities

| Variable | (1) Model 5 | (2) Model 6 | (3) Model 8 |
|---|---|---|---|
| FLTEAR | -4.002 | -3.791 | -3.689 |
| FLTEAR2 | 0.157 | 0.190 | 0.206 |
| RACE | 0.977 | 1.097 | 1.059 |
| SERLEN | 0.008 | 0.007 | 0.005 |
| _SERLEN | -0.042 | -0.030 | -0.010 |
| SERLEN2 | 0.000 | 0.000 | -0.001 |
| _SERLEN2 | 0.001 | 0.001 | 0.000 |
| RAGER1 | -0.044 | 0.117 | 0.007 |
| RAGER2 | -0.528 | -0.565 | -0.299 |
| RAGER3 | 0.849 | 1.167 | 3.240 |
| _RAGER1 | -0.546 | -0.435 | -0.270 |
| _RAGER2 | -0.295 | -0.297 | -0.505 |
| RCOHORT1 | 1.788 | 0.929 | 0.844 |
| RCOHORT2 | 1.444 | 0.891 | 0.817 |
| _RCOHORT1 | 5.359 | 3.711 | 2.401 |
| _RCOHORT2 | 3.536 | 2.357 | 1.098 |
| EDU1 | -0.301 | -0.270 | -0.355 |
| EDU2 | 0.084 | 0.151 | 0.050 |
| EDU3 | 0.091 | 0.159 | -0.066 |
| EDU4 | 0.969 | 1.054 | 0.998 |
| _EDU1 | -0.038 | -0.071 | -0.529 |
| _EDU2 | 0.309 | 0.239 | 0.345 |
| _EDU3 | 0.055 | 0.064 | 0.060 |
| _EDU4 | 0.055 | 0.009 | 0.469 |
| Intercept | 92.520 | 94.167 | 95.230 |
| $R^2$ | 0.846 | 0.845 | 0.612 |

However, the coefficient for RAGER3 becomes progressively larger as the annuity counterfactual more closely replicates the OAI program. Looking at the comparable variables for men, the coefficient for _RAGER1 decreases in size, whereas the coefficient for _RAGER2 increases in size as additional program features are included in the annuity counterfactual. Third, the coefficients for the retirement cohort variables (RCOHORT1, RCOHORT2, _RCOHORT1, _RCOHORT2) systematically decrease across the model variations.

Similar results, although not identical measures, are observed in Table 7.36.

3. <u>Summary of regression findings</u>

a. <u>Lifetime earning variables (LTEAR, LTEAR2, FLTEAR, FLTEAR2)</u>
For all permutations of the single and married models, the estimated coefficient for the household measure of lifetime earnings was negative. This suggests that, when all other household characteristics were held constant, households with higher lifetime earnings received smaller redistribution components. The relationship between percentage of redistribution and lifetime earnings was, however, nonlinear (the coefficient $B_1$ is negative and $B_2$ is positive). Thus, the percentage of redistribution decreases at a decreasing rate as lifetime earnings increases. (Technically, the percentage of redistribution will at first decrease but later increase as lifetime earnings increases; however, given the range of LTEAR and FLTEAR in this study, the measured

relationship between the percentage of redistribution and lifetime earnings was negative.)

The inclusion of disaggregated survivorship probabilities did not reverse the relationship between the redistribution measure and lifetime earnings. For the single model, accounting for mortality differentials by sex and race generally weakened the relationship between the redistribution and earnings measures. But, further disaggregation by marital status, income, and education tended to strengthen the relationship over comparable estimates using sex and race differentials and, in several cases, over the similar estimates for age-only mortality differentials. On the other hand, for married households, the relationship between the redistribution measures and lifetime earnings was consistently weakened when sex-race and sex-race-marital status-income-education differentials were introduced. Therefore, it is not accurate to conclude that highly disaggregated mortality rates reverse or substantially weaken the progressivity of the program. From the findings on married households, mortality rates disaggregated by sex and race challenge the program's progressivity slightly less than mortality rates disaggregated by sex, race, marital status, income, and education.

Findings on the effect of benefit indexing on the relationship between the percentage of redistribution and lifetime earnings are consistent across household types. For single households, the inclusion of indexing in the annuity counterfactual slightly weakens the negative relationship between the percentage of redistribution and lifetime earnings in the models with the earnings. That is, independent of the degree of mortality rate disaggregation, the OAI program was found to be less progressive for single households after the inclusion of benefit indexing in the annuity counterfactual. For all the married models, there is a stronger negative relationship between the percentage of redistribution of lifetime earnings without indexing. Hence, in all of the single married models, the OAI program is less progressive when the annuity counterfactual includes benefit indexing.

The addition of the earnings test consistently weakens the relationship between the percentage of redistribution and lifetime earnings for married households, but it consistently strengthens the relationship for single households. These findings are suggestive of different employment decisions by single and married households after retirement. The strengthening of the relationship for single households would seem to indicate that single persons with lower lifetime earnings were more inclined to work after retirement. After examining the data set, it was found that 25 single households were affected by the earnings test, of which 65 percent were women. A majority of the households affected by the earnings test had lifetime earnings measures below the sample average. The labor force attachment of women after retirement may reflect not only the sex distribution of the retirement population, but that single women typically have less physically demanding occupations which characteristically permit greater staying power. The opposite was true for the married population. Generally speaking, high income, married persons tended to continue working after retirement. The employment pattern of the married households is consistent with studies on retirement patterns (Boskin, 1977, Pechman et al., 1968). In 1966, only 1.6 million of the 17 million persons eligible for retirement benefits were affected by the retirement test. Fifty percent of the 1.6 million beneficiaries affected by the earnings test earned $2,700 or more in 1966.

In general, the OAI program was found to be progressive with respect to lifetime earnings across all model permutations. The strength of the negative association between household percentage of redistribution and household earnings varied by marital status. In particular, the program had stronger progressive features for single households relative to married households. This finding is not too surprising in light of the extra benefits extended to wives of covered workers.

b. Service length (SERLEN, SERLEN2) For the single models, the coefficients on SERLEN and SERLEN2 are negative and positive, respectively, and the coefficients for SERLEN are significantly different from zero at a one percent level, but the coefficient for SERLEN2 is statistically significant in model 2 only. The coefficient for SERLEN is remarkably stable across the models, whereas the estimated coefficient for SERLEN2 modestly increases when program features are added to the annuity counterfactuals. The estimated negative, nonlinear association between the percentage of redistribution and service length suggests that longer contribution periods significantly reduce the percentage of redistribution received in retirement, *ceteris paribus*.

The comparable sex-coded estimates for married persons are mixed and statistically insignificant. Again, the estimated coefficients are remarkably stable across survivorship assumptions. However, the coefficients for the service length variables SERLEN and _SERLEN tend toward zero when additional program features are introduced into the annuity counterfactual.

c. Sex (SEX) The sex variable was included in the single model only. Without mortality differentials by sex, benefit indexing, and the earnings test, the coefficient for SEX is negative and statistically insignificant. However, with the inclusion of sex differentials in survivorship, the coefficient for SEX is positive and significantly different from zero at a one percent level. Further disaggregation of mortality differentials by marital status, income, and education increased the redistributional gains of single women over single men. As a result of their longevity, women received significant redistributional gains from the OAI program, *ceteris paribus*.

Single women, also, received further redistributional gains when benefit indexing and survivorship differentials by sex were included in the annuity counterfactual. The addition of the earnings test did not appreciably affect the female-to-male difference in the percentage of redistribution after accounting for benefit indexing and mortality differentials. Overall, females received redistribution components approximately five percentage points larger than their male counterparts when indexing, post-retirement earnings adjustments, and mortality differentials by sex were reflected in the annuity counterfactual.

d. Race (RACE) In the case of single households, the coefficient for RACE is small and statistically insignificant, with the notable exception of models 2 and 4 when benefit indexing and mortality differentials by sex, race, marital status, income, and education were accounted for in the annuity counterfactual. The sign for the RACE coefficient is mixed and dependent on counterfactual characteristics. Under the assumptions of model 1, the coefficient for RACE is positive, suggesting that nonwhites received a redistribution component slightly larger than their white counterparts, *ceteris paribus*. The slight gains of nonwhites are probably symptomatic of earnings differentials by race prevalent in the labor market. Whites, on average, receive higher earnings relative to nonwhites, concentrating nonwhites at the lower end of the progressive benefit formula. Even with the adjustments for race differentials in mortality, the nonwhite redistributional gain persisted. This seems to suggest that OAI benefit differentials by race were stronger than mortality differentials by race. Examining the survivorship probabilities by race and sex in Table 13.5, it is observed that mortality differentials by race are fairly weak and the sign of the differential reverses at advanced ages. Generally speaking, however, disaggregated mortality rates reduced the size of the nonwhite gain, and oft-times resulted in redistributional losses.

The inclusion of benefit indexing in the annuity counterfactual and mortality differentials by race and sex result in estimated coefficients that are negative. Further disaggregation of mortality rates by marital status, income, and education, result in estimated coefficients that are negative and statistically significant. Identical results occur with the addition of the earnings test.

Looking at the married models, the coefficients for RACE are negative and statistically insignificant except when mortality differentials are disaggregated by sex, race, marital status, income, and education. The negative relationship between race and household percentage of redistribution is strengthened when mortality differentials by sex and race are included; however, when mortality differentials by sex, race, marital status, income, and education are included, the estimated coefficient for race is positive and statistically significant except in model 7. The effect of benefit indexing and the earning test features on the coefficient estimate is dependent on the survivorship assumption; using the gender-merged and sex-race survivorship probabilities, the race differential is weakened with indexing but strengthened with the earning test; however, using the socioeconomic-adjusted probabilities, the race differential is strengthened with indexing but weakened with the earnings test.

The mixed and contradictory results across married models and across the married and single models are perplexing. One contributing factor for the erratic performance of the race variable is the weak representation of nonwhites in the data set. Nonwhites accounted for six percent of the single households and two percent of married households. Clearly, any generalizations based on the size and sign of the estimated coefficients for RACE are tenuous and should not be taken too seriously.

e. Age at retirement (RAGER1, RAGER2, RAGER3) Most of the evidence on the age of retirement suggests that single persons received the largest redistributional component by retiring at age 65, *ceteris* paribus. This finding is consistent with earlier mentioned criticisms of the actuarial adjustment formulas.

Looking first at RAGER1, the variable for retirement prior to age 65, the coefficient for RAGER1 using the gender-merged survivorship table is negative and statistically insignificant across all versions of the generalized single model. The inclusion of disaggregated mortality differentials reduces the size of the negative redistribution differential for persons who retired earlier than age 65, and, in some cases, reverses the sign of the redistribution differential. The addition of the benefit indexing feature to the annuity counterfactuals reverses the sign of the coefficient for RAGER1, whereas the earnings test feature does not significantly affect the size or sign of the coefficient.

The coefficient for RAGER2 is negative and statistically insignificant for all permutations of the generalized single model. The strength of the negative relationship decreases as mortality differentials are disaggregated. Similarly, benefit indexing and earnings test provisions further weaken the difference between the redistribution differential for persons retiring between ages 65 and 71 relative to persons retiring at age 65, ceteris paribus.

The last age at retirement variable to be discussed is RAGER3. The coefficient for RAGER3 is negative and significantly different from zero at a one percent level for all single models. The size of the redistributional differential is augmented by mortality rate disaggregation, benefit indexing, and earnings test adjustments, with the notable exception of model 4 using mortality differentials by sex and race.

Next, looking at the sex-coded age at retirement variables for the married model, the results for RAGER1 are mixed and statistically insignificant. Early retirement for women does not significantly affect the size of the household redistribution measure relative to households where the woman retired at age 65, _ceteris paribus_. However, the household redistribution measure is slightly smaller when the woman retired between the ages of 65 and 71 relative to age 65, _ceteris paribus_. The size of loss is slightly increased with increased disaggregation of mortality rates and the introduction of benefit indexing, but it is slightly reduced with the earnings test. The last age of retirement variable is RAGER3. The coefficient for RAGER3 is positive and significantly different from zero at a one percent level for all models. The strength of the positive relationship is augmented by mortality rate disaggregation, benefit indexing, and earnings test adjustments. It is not surprising that women who postponed retirement to age 72 or later received abnormally high household redistribution measures. These women were most probably collecting special age-72 benefits, which are provided to aged persons who cannot claim benefits as a primary worker or dependent spouse and who have very few quarters of coverage; hence, OAI benefits were received by these women at a near-zero cost.

The coefficients for _RAGER1 and _RAGER2 are negative and significantly different from zero at a one percent level. The strength of the relationship is weakened by mortality disaggregation by sex and race, but largely unaffected by further socioeconomic disaggregation. The household redistribution differential for males who retired after (before) age 65 increased (decreased) in magnitude with the inclusion of benefit indexing and earnings test adjustments in the annuity counterfactual.

f. Retirement cohort (RCOHORT1, RCOHORT2) Estimates of the coefficients for RCOHORT1 and RCOHORT2 are positive and significantly different from zero at a one percent level for all permutations of the single and married generalized models. Also, the size of the coefficient for RCOHORT1 exceeds the size of the coefficient for RCOHORT2, suggesting that the gains from retiring in an earlier retirement cohort diminish over time.

For the single models, the effect of disaggregated mortality rates are mixed. When mortality differentials disaggregated by sex and race were used, the estimated coefficients for RCOHORT1 and RCOHORT2 diminish in size, reducing the intercohort redistributional differential. However, further disaggregation places upward pressure on the estimated size of the RCOHORT1 and RCOHORT2 coefficients; hence, the intercohort redistributional differential widens. It appears that the earlier cohorte had different educational and income characteristics which tended to reverse the influence of sex and race differentials in survivorship on the redistribution measure.

The addition of benefits indexing and the earnings test to the annuity counterfactual systematically narrows the intercohort redistributional differential, as expected. Since this study evaluates the OAI program in 1972 and retirement cohorts from 1962 to 1972 are included in the data set, benefit levels promised in real terms must be augmented over the retirement interval from 1962 to the year of program assessment, 1972. The benefit adjustment scheme indexed the initial annuity benefit in the retirement year by $(1+c)^t$, where c equals .0275 (the annuitized rate for future price changes) and t equals the difference between the retirement year and 1972. Because of _ex post_ indexing, the intercohort redistributional differential is narrowed. The narrowing effect of the earnings test feature was also expected since the 1969-1972 retirement cohort had the greatest likelihood of receiving labor earnings in excess of the earnings limit in 1972, which would place upward pressure on the size of later cohorts' redistribution components, subsequently narrowing the redistributional differential across cohorts.

For the married models, similar results are obtained for the female-coded RCOHORT1 and RCOHORT2 coefficients. That is, disaggregated mortality differentials, benefit indexing, and the earnings test adjustments tended to narrow the intercohort redistributional differential. However, the male-coded _RCOHORT1 and _RCOHORT2 coefficients are invariant to the level of mortality rate disaggregation, but they tended to diminish in size with the addition of benefit indexing and the earnings test, _ceteris paribus_.

g. _Level of education (EDU1, EDU2, EDU3, EDU4)_ With the exception of the coefficient for EDU4, the estimated coefficients for the education variables in the single models are generally negative and statistically insignificant. That is, the redistributional differential by education level is negative albeit small for households with less than eight years of education or high school training relative to households with eighth grade education only. The influence of different mortality rate assumptions are mixed. For households with less than eight years of education or more than 12 years of education, the inclusion of sex and race differentials in survivorship tended to either eliminate existing redistributional gains or increase redistributional losses relative to households with eighth grade educations. However, further disaggregation of mortality rates by marital status, income, and education generally reduced the redistributional gap between households with eight years of education and those with 12 or more years of education, but expanded the gap between households with eight years of education and those with less than eight years. This result is reflective of the inverse relationship between mortality and education and income levels. Mortality disaggregation tended to eliminate the negative differential between households with 9-11 years of education and eight years of education. Furthermore, benefit indexing narrowed the education redistributional differential. But, the earnings test tended to widen the differential for households with 0-7 and 12 years of education, while it narrowed the differential for households with 9-11 and 13 or more years of education. The earnings test effect suggests that persons with 9-11 or 13 or more years of education tended to remain in the labor force after retirement.

Again, the coefficient estimates for the sex-coded education variables are mixed and generally statistically insignificant. However, a few general patterns are worth mentioning. For all education groupings excluding EDU4, the inclusion of sex and race differentials in survivorship tended to narrow the education redistributional differentials, whereas further disaggregation tended to improve the redistribution status of households with any of the following education variables: EDU2, EDU3, EDU4, _EDU3, and _EDU4. The inclusion of the earnings test greatly increased the positive redistribution differential for males with college education, while it increased the negative redistributional differential for households with any of the following education variables: EDU1, EDU2, EDU3, and _EDU1. Again, these results are reflective of post-retirement employment patterns of married households.

# VIII. SUMMARY AND CONCLUSIONS

## A. Summary

Chapter I presented a brief overview of the federal old-age insurance program and the method employed to isolate the distributional impact of the social security program. Also, the four interrelated issues investigated in this study were identified.

Chapter II was a detailed discussion of the historical development of the OAI program with emphasis on the following program features: spousal benefits, progressive benefit formula, actuarial reduction for early retirement, delayed retirement credit, earning test formula, and cost-of-living adjustments. Each program feature was explained in terms of its original intent, redistributive effect, and controversial implications, when applicable.

Previous empirical studies on the distributional impact of the social security program were reviewed in Chapter III. Virtually all empirical studies indicate that social security beneficiaries retiring prior to 1975 received above-normal rates of return on their contribution dollars; however, there was less agreement concerning the program's progressivity. Empirical evidence did support allegations that the intent of many program features were compromised by the program's design and demographic characteristics of the retirement population. While the cited studies differed in detail, the distributional impact measure (be it an internal rate of return, contribution-benefit ratio, or transfer component) was found to be sensitive to specific identifiable worker characteristics, such as date of retirement, marital status, sex, race, income, education level, and age at entry and retirement. The distributional significance of each worker characteristic was discussed in Chapter III.

A life-cycle model for evaluating the distributional impact of the OAI program was presented in Chapter IV. Two conditions for an actuarially fair retirement program were specified, which were subsequently used to explain the "disentanglement" of OAI benefits along functional lines.

Chapter V describes the methodology. The model assumptions regarding the fairness standard, interest rates, survivorship probabilities, earnings test formula, and behavioral responses were discussed in detail. Also, a description of the data set, computational formulas, annuity-type counterfactuals, and redistribution components were presented.

The generalized polynomial regression models by marital status were described in Chapter VI. A generalized model was specified for the purpose of drawing inferences regarding the effect of worker and program characteristics on the distributional impact of the OAI program. The dependent and independent variables were defined and described in Chapter VI.

Descriptive statistics, in tabular array, on the benefit incidence for all households, single households, and married, both retired households and the results of the empirical analysis of this study were presented in Chapter VII. The descriptive statistics indicated that:

1) all family types received more than their "money's worth" from the OAI program in 1972; 2) single females and married couples were made better off, and single males were made worse off in a sex-neutral retirement program; 3) traditional family structures received preferential treatment from the OAI program because the dependent spouse received retirement benefits without payment of extra contributions; 4) dependent's benefits were equally distributed across quintile groups; and 5) the OAI program tended to be more regressive with the introduction of the earnings test and socioeconomic-adjusted survivorship rates.

Evidence from, and interpretation of, the ordinary least-square multiple regression estimation of the polynomial models was presented in Chapter VII. The regression estimates did, in most cases, support the generalizations derived from the descriptive statistics.

## B. Conclusions

Four interrelated issues were addressed in this study: 1) Does the OAI portion of the social security program redistribute income in favor of low-income beneficiaries? 2) Does the current OAI program redistribute benefits in favor of women, as a group, at the expense of their male counterparts? 3) How does the wife's work status affect the distribution of OAI benefits within and across family types? 4) Are spousal benefits distributed principally to needy dependent spouses? Answers to these questions will be presented in this final section.

1. Overall program assessment

The results presented in Chapter VII suggest that for OAI beneficiaries the program was progressive with respect to income. Thus, the program did tend to favor low-income beneficiaries in terms of the percentage of redistribution. Tabular results showed that all income groups received more than their "money's worth" from the OAI program in 1972; however, the largest relative gains were realized by low-income groups. Using different program assessment approaches, the OAI program in 1972 for the full data set was found to be "mildly" and "generally" progressive across income groups, but it also exhibited strong regressive features, resulting in lower relative returns to middle-income beneficiaries.

The program was found to be more effective in redistributing income in the absence of the earnings test, price indexing, and disaggregated survivorship probabilities. First, the earnings test, in general, had its greatest impact on high-income families ($6,000+), which tended to increase the percentage of redistribution received by high-income families. According to the design of this study, the OAI program was found to be more regressive after the earnings test feature was included into the analysis. Second, at least initially, all income classes received larger redistribution components when the annuity counterfactual was defined to include price indexing. Although all income groups realized extra redistribution per dollar of OAI benefits when indexing was included in the analysis, the greatest relative gains were realized by higher income groups because of their longer life expectancies on average. Price indexing, itself, does not alter the redistribution pattern, although it does slightly affect the levels of progressivity and regressivity as measured by the "high-income-group-comparison" approach to progressivity assessment. In addition to the aforementioned program features, demographic factors, such as differential survivorship probabilities, do have an unintended effect on the equity of the program. Based on tabular results for the full data set, the program's overall progressivity was found to be virtually invariant to the use of gender-merged or sex-race-distinct survivorship probabilities; however, slight progressivity changes were observed with the use of socioeconomic-adjusted survivorship probabilities. Specifically, the program had slightly weakened progressive features for low-income households and slightly strengthened regressive features for middle-income households when socioeconomic differentials in survivorship were incorporated into the counterfactual design.

The tabular results for the full data set are largely supported by the regression results. However, two cautionary notes should be mentioned regarding any direct comparisons between tabular and regression findings. First, the tabular and regression findings are based on different groupings of the same retirement population. That is, the tabular results discussed above were based on the full data set including single households and married households where at least one member of the couple was retired in 1972. On the other hand, the regression results are based either on the single household or married households where both members were retired in 1972. Because of the different groupings, the results may appear to be contradictory when taken together, although, when taken separately, they are consistent with a priori reasoning. Second, the measures of income are different for the tabular and regression analyses. The tabular results were arrayed by family income in 1972, as reported on the 1973 census questionnaire. The regression results are based on lifetime earnings, a summary statistic representing the accumulated value of annual taxable real earnings for the household unit. Each earnings measure has obvious shortcomings and was used to achieve different ends. The tabular results are directly comparable with Burkhauser and Warlick's (1981) presentation, whereas the regression results are directly comparable with Freiden et al. (1976). Hence, the earnings measures, while complicating comparisons within the study, are perfectly useful across previous studies.

The regression results support the findings of the program's progressivity. Recall, the estimated relationship between the percentage of redistribution and lifetime earnings (LTEAR or FLTEAR) was negative and nonlinear. The effect of the earnings test was mixed and dependent on marital status. The progressivity of the program was weakened for the married data set and strengthened for the single data set with the inclusion of the earnings test in the annuity counterfactual. Evidently, the "married" influence of the earnings test dominated when the data were aggregated in the tabular results. Similarly, in all of the single married models, the OAI program was less progressive when the counterfactual included benefit indexing. The inclusion of disaggregated survivorship probabilities did not reverse the relationship between the percentage of redistribution and lifetime earnings, and, in addition, the marginal effect, overall, was small. From the regression findings, mortality rates disaggregated by sex and race challenge the program's progressivity slightly less than highly disaggregated mortality rates. Contrary to Aaron's study (1974), the effect of socioeconomic differentials in survivorship does not reverse the direction of redistribution, but, rather, slightly "dampens" the extent of redistribution.

## 2. The effect of sex differentials in survivorship on the program's performance

The distributional impact of the OAI program was found to be sensitive to the "tailoring" of annuity benefits to reflect sex differentials in survivorship. Generally speaking, single females and married couples were made differentially better off, and single males worse off in a sex-neutral retirement system relative to a sex-discriminating actuarially fair retirement system. Single female beneficiaries, as a group, received annuity benefits that were approximately 16 percent larger in a sex-neutral retirement system relative to a sex-race discriminating system, whereas their male counterparts, as a group, received benefits that were approximately seven percent smaller. Furthermore, when the mortality differentials were disaggregated by sex, race, marital status, income, and education, single female beneficiaries received annuity benefits that were approximately nine percent larger in a sex-neutral retirement system, whereas single male beneficiaries received benefits that were approximately 23 percent smaller.

Similar comparisons were not as useful across married beneficiaries because the joint-and-two-thirds annuity covered the lives of the husband and wife; hence, any sex differentials were largely muted by the dual coverage. Nonetheless, actuarially fair benefits for married persons were approximately three percent higher, independent of the sex of the primary annuitant, in a sex-neutral retirement system relative to a sex discriminating system. The sex-neutral bias in favor of married persons, as a group, is a result of the joint-and-two-thirds annuity, which insures the life of the shorter-lived male, the longer-lived female, and the longest-lived survivor, who is typically the female. The sex-neutral bias increased when the socioeconomic discriminating system was used as the comparison system.

The estimated coefficient for SEX in the single generalized model was positive and statistically significant, supporting the tabular findings. Single female beneficiaries received redistribution components approximately three percentage points larger than their male counterparts when survivorship probabilities were disaggregated by sex and race, *ceteris paribus*. The marginal gain increased to 5.2 percentage points when survivorship probabilities were further disaggregated by marital status, income, and education.

3. <u>The effect of the wife's work status on the program performance</u>

The influence of the wife's work status was examined extensively in section A of Chapter VII. To address this issue, households, where both members were retired in 1972, were divided into one-earner and two-earner households. A two-earner household was defined as a household where both members qualified for primary-worker benefits. Alternatively, a one-earner household meant only the male member qualified for primary-worker benefits and the spouse was collecting dependent's benefits. Independent of sex and family type, all individuals received positive income transfers from the OAI program in 1972. Overall, the traditional family structure received preferential treatment from the OAI program because the nonworking wife received retirement benefits without payment of extra contributions.

First, the effect of the wife's work status on wife-only benefit incidence was small. In absolute terms, working women paid in more dollars in the form of OAI contributions, and, in exchange, they received higher OAI benefit levels. However, the difference in percentage of redistribution per dollar of OAI benefits for working and nonworking women was extremely small, suggesting that women, independent of work status, were treated almost equally in terms of redistribution.

The finding of roughly equal treatment across women with different labor-homemaker choices did not apply to men married to women with different labor-homemaker choices. Generally speaking, the percentage of redistribution was generally higher for males in one-earner households relative to their male counterparts in two-earner households. The apparent redistributional differential was symptomatic of the very low annuity benefits received from the nonworking wife's joint-and-two-thirds annuity.

In conclusion, although women with different work statuses paid in different amounts of OAI contributions, they were treated roughly equally in terms of the percentage of OAI benefits representing intergenerational transfers. The redistribution pattern for males by household type was similar; however, the absolute size of the percentage of redistribution was larger for one-earner males across all income categories. While women were treated roughly equally, working women received significantly smaller percentage of redistribution when comparisons were made with working males. The working woman received the smallest return on her OAI contributions relative to her male counterpart because of her retirement and employment characteristics and the community property assumption underpinning the annuity-type counterfactual. Lastly, the OAI program was found to be more progressive and less regressive across income categories for two-earner relative to one-earner households as reflected by the "high-income-group-comparison" approach to progressivity assessment.

4. The distribution of spousal benefits

The OAI program was found to allocate redistribution components proportionately across quintile groups, independent of family type and sex. Contrary to the 1937-1939 Advisory Council's intent, dependent's benefits were, at best, proportionally distributed to dependent spouses of male workers. Twenty-two of the poorest one-earner households received approximately 21 percent of intergenerational transfers to dependent spouses compared to 21 percent received by the 20 percent of the richest one-earner households. Evidence from this life-cycle study supports the earlier findings of Holden (1979). In conclusion, this study demonstrated that supplemental benefits may not be adequately serving the 1939 objective of protecting a group of aged persons experiencing economic hardship, suggesting, perhaps, that a more effective target definition should be used to determine "need" aside from the work status of the female, which is currently used by social security.

## C. Concluding Remarks

This study attempted to estimate the extent to which the old-age insurance portion of the social security program redistributed income among subgroups comprising the same retirement population but distinguishable by socioeconomic traits, such as sex, race, marital status, income, and education. In estimating the distributional impact of the social security program, the study stressed the importance of an intertemporal framework to evaluate a "lifetime" public program and the need to account for demographic factors, such as differential mortality rates.

Overall, the 1972 OAI progam was found to be progressive; however, "other" program features and socioeconomic status were also found to influence the effectiveness of the program in achieving its redistribution objective. From a policy point of view, this study has several noteworthy implications. First, evidence from this study showed that the OAI program, as legislated in 1972, was not distributionally neutral, and its distributional impact oft-times depended on factors incidental to the program. Second, the legislated preferential treatment of women, traditional family structures, and earlier retirement cohorts draw into question and challenge the redistribution objective of the OAI program. Third, it was found that the intended and actual effects of statutory provisions may vary widely and may, as a result, jeopardize the effectiveness of the program in general. In the future, policymakers should be cognizant not only of the intended and actual effects of statutory provisions, but also of the unintended effects of demographic factors, incidental to the program, on the overall equity of the social security system.

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# XI. APPENDIX A. ESTIMATION OF MARKET YIELDS ON U.S. GOVERNMENT SECURITIES AT CONSTANT MATURITY, 1937-1952

The roll-over compounding scheme was estimated using a historical series of market yields on U.S. government securities at constant maturity. Yields are reported on 1, 3, 5, 10, 20 and 30 year maturities for the 1953 to 1972 time period (Board of Governors, 1976b); however, disaggregated data were not available for earlier years. For the years prior to 1953, the missing yields were estimated using known yields on 3-5 year taxable U.S. notes, 1937-1970 (Board of Governors, 1943; U.S. President, 1976), and a historically complete series of basic yields on corporate bonds by term to maturity (Board of Governors, 1943, 1976a). The private bond yield curve for each maturity in conjunction with the 3-5 year taxable note series were used to replicate the yield curves for U.S. government securities for the missing years. The procedure employed to complete the series is discussed in detail below.

Firstly, the basic yields on corporate bonds are reported at one-year maturity intervals for corporate bonds with ten or fewer years to maturity. A 3-5 year yield series for prime corporate paper for 1937 to 1970 was constructed by taking an arithmetic average of the three-year and five-year yields for each year between 1937 and 1970. The 3-5 year yield for each year, 3-5 PMY, is represented by

$$\text{3-5 PMY}_y = \frac{\text{3-year PMY}_y + \text{5-year PMY}_y}{2} \tag{11.1}$$

where $PMY_y$ = prime market yield in year y. The 3-5 $PMY_y$ was used as a standard of comparison to simulate the yield curves for U.S. government securities at constant maturity for the missing years.

Next, yields on U.S. government securities at constant maturities were calculated as follows:

$$GMY_y^i = (3\text{-}5\ GMY_y) \times \frac{PMY_y^i}{3\text{-}5\ PMY_y} \tag{11.2}$$

where $GMY_y^i$ = estimated U.S. government security market yield at maturity i in year y,

i = years to maturity (1, 3, 5, 10, 20, 30),

3-5 $GMY_y$ = yield for 3-5 year taxable U.S. securities for year y, and

$PMY_y^i$ = yield on prime corporate bond at maturity i in year y.

The estimated market yields are shown in Table 11.1.

The accuracy of the above estimation procedure was tested by comparing the known U.S. government security yields to the estimated yields for the 1953-1972 time period. Comparisons are shown for the 5, 10, and 20 year maturities in Table 11.2. The size of the estimation error is less than five percent for most maturities and years. Estimation errors are largest in years 1954, 1958, and 1959. However, the estimation error is smaller than the error resulting from the use of the prime corporate bond yield in place of the U.S. government security rate (see Figure 11.1).

Table 11.1. Market yields on U.S. government securities at constant maturity, 1937-1972 (percent per annum)

| Year | 1-year | 3-year | 5-year | 10-year | 20-year | 30-year |
|---|---|---|---|---|---|---|
| 1937 | 0.66 | 1.19 | 1.60 | 2.27 | 2.76 | 2.94 |
| 1938 | 0.41 | 0.72 | 0.94 | 1.24 | 1.39 | 1.43 |
| 1939 | 0.26 | 0.50 | 0.93 | 0.96 | 1.17 | 1.21 |
| 1940 | 0.19 | 0.42 | 0.58 | 0.90 | 1.17 | 1.24 |
| 1941 | 0.29 | 0.61 | 0.85 | 1.34 | 1.77 | 1.88 |
| 1942 | 0.875 | 1.33 | 1.59 | 1.93 | 2.46 | 2.46 |
| 1943 | 0.75 | 1.22 | 1.45 | 1.96 | 2.47 | 2.47 |
| 1944 | 0.79 | 1.21 | 1.44 | 1.94 | 2.48 | 2.48 |
| 1945 | 0.81 | 1.07 | 1.29 | 1.60 | 2.37 | 2.39 |
| 1946 | 0.82 | 1.04 | 1.26 | 1.80 | 2.25 | |
| 1947 | 0.92 | 1.19 | 1.45 | 1.83 | 2.11 | |
| 1948 | 1.34 | 1.54 | 1.72 | 2.14 | 2.30 | |
| 1949 | 1.24 | 1.37 | 1.49 | 1.8 | 2.03 | |
| 1950 | 1.2 | 1.40 | 1.61 | 1.94 | 2.09 | |
| 1951 | 1.81 | 1.89 | 1.97 | 2.12 | 2.30 | |
| 1952 | 2.13 | 2.13 | 2.13 | 2.13 | 2.24 | |
| 1953[a] | 2.14 | 2.47 | 2.65 | 2.85 | 3.06 | |
| 1954 | 1.05 | 1.63 | 1.99 | 2.40 | 2.64 | |
| 1955 | 2.04 | 2.47 | 2.65 | 2.82 | 2.90 | |
| 1956 | 2.99 | 3.19 | 3.20 | 3.18 | 3.14 | |
| 1957 | 3.62 | 3.98 | 3.69 | 3.65 | 3.54 | |
| 1958 | 2.27 | 2.84 | 3.06 | 3.32 | 3.48 | |
| 1959 | 4.24 | 4.46 | 4.46 | 4.33 | 4.13 | |
| 1960 | 3.63 | 3.98 | 4.09 | 4.12 | 4.06 | |
| 1961 | 2.98 | 3.54 | 3.75 | 3.88 | 3.92 | |
| 1962 | 3.10 | 3.47 | 3.70 | 3.95 | 3.99 | |
| 1963 | 3.36 | 3.67 | 3.83 | 4.00 | 4.05 | |
| 1964 | 3.85 | 4.03 | 4.07 | 4.19 | 4.19 | |
| 1965 | 4.14 | 4.22 | 4.25 | 4.28 | 4.27 | |
| 1966 | 5.20 | 5.23 | 5.10 | 4.92 | 4.77 | |
| 1967 | 4.88 | 5.03 | 5.11 | 5.07 | 5.01 | |
| 1968 | 5.69 | 5.68 | 5.69 | 5.65 | 5.45 | |
| 1969 | 7.12 | 7.02 | 6.93 | 6.67 | 6.33 | |
| 1970 | 6.90 | 7.29 | 7.38 | 7.35 | 6.86 | |
| 1971 | 4.88 | 5.65 | 5.99 | 6.16 | 6.12 | |
| 1972 | 4.96 | 5.72 | 5.98 | 6.21 | 6.01 | |

[a]Board of Governors, 1976b.

Table 11.2. Estimation error

| Year | 5-year estimation error | | 10-year estimation error | | 20-year estimation error | |
|---|---|---|---|---|---|---|
| 1953 | 0.06 | (2.3)[a] | 0.14 | (4.9)[a] | 0.19 | (6.2)[a] |
| 1954 | 0.15 | (7.5) | 0.45 | (19.0) | 0.53 | (20.0) |
| 1955 | 0.12 | (4.5) | 0.22 | (7.8) | 0.15 | (5.2) |
| 1956 | 0.01 | (0.3) | -0.11 | (3.5) | -0.27 | (8.6) |
| 1957 | 0.07 | (1.9) | 0.03 | (0.8) | -0.08 | (2.3) |
| 1958 | 0.16 | (5.2) | 0.33 | (10.0) | 0.38 | (10.9) |
| 1959 | 0.09 | (2.0) | -0.30 | (6.9) | -0.59 | (14.3) |
| 1960 | 0.14 | (3.4) | 0.29 | (7.0) | 0.27 | (6.7) |
| 1961 | 0.01 | (0.3) | -0.12 | (3.0) | -0.22 | (5.6) |
| 1962 | 0.02 | (0.5) | -0.01 | (0.3) | -0.08 | (2.0) |
| 1963 | -0.04 | (1.0) | -0.06 | (1.5) | -0.15 | (3.7) |
| 1964 | -0.03 | (0.7) | -0.01 | (0.2) | -0.07 | (1.7) |
| 1965 | -0.01 | (0.2) | -0.02 | (0.5) | -0.03 | (0.7) |
| 1966 | -0.01 | (0.2) | -0.19 | (3.9) | -0.18 | (3.8) |
| 1967 | 0.04 | (0.8) | 0.05 | (1.0) | 0.19 | (3.8) |
| 1968 | 0.1 | (1.8) | 0.12 | (2.1) | 0.08 | (1.5) |
| 1969 | 0.08 | (1.1) | -0.18 | (2.7) | -0.25 | (3.9) |
| 1970 | -0.01 | (0.1) | 0.05 | (0.7) | -0.07 | (1.0) |

[a]Error as a percentage of the known yield to maturity on U.S. government securities in each year.

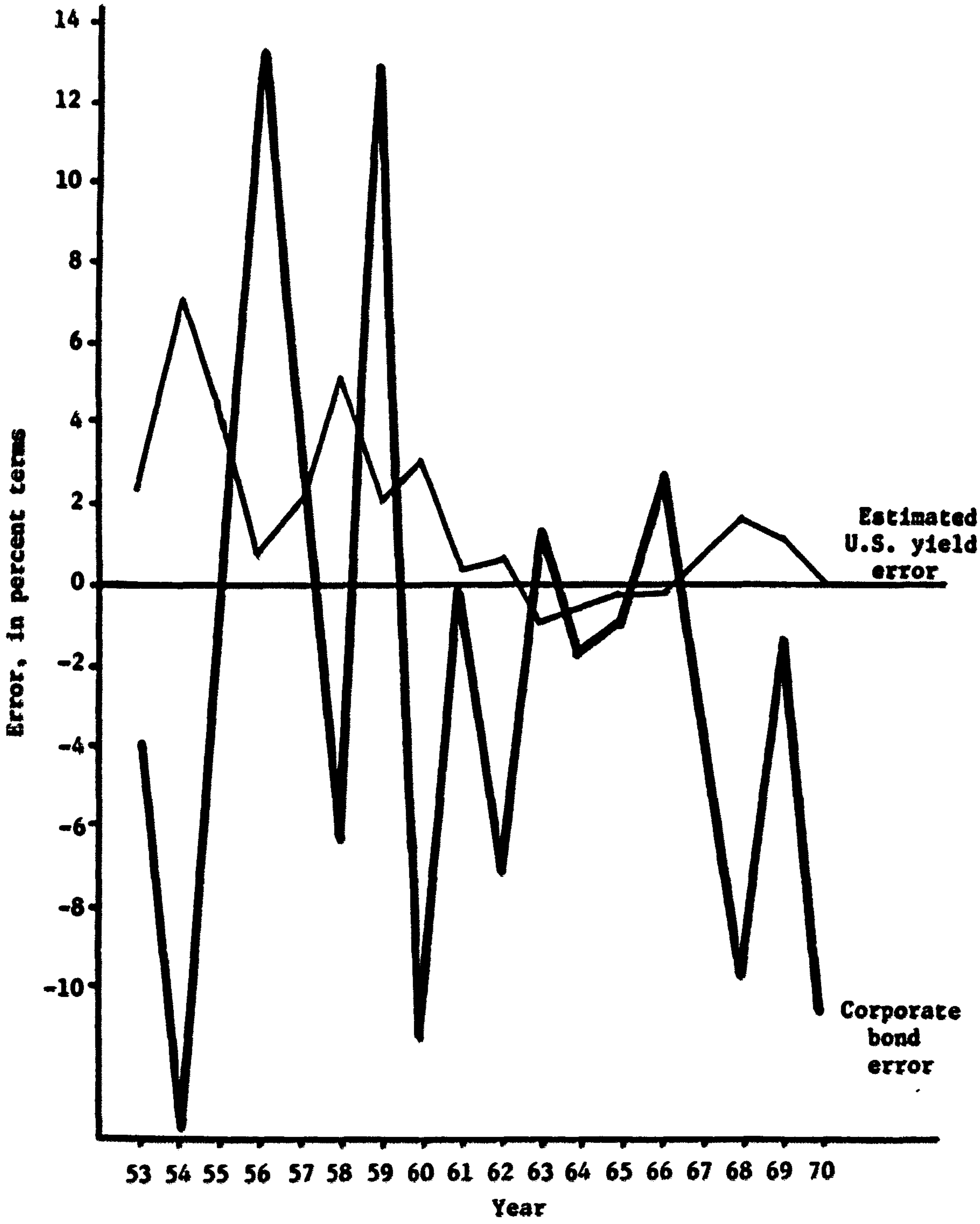


Figure 11.1. Comparison in five-year estimation errors: Estimated U.S. yields relative to known yields for 1953-1970 and corporate bond yields relative to yields on U.S. security yields for 1953-1970

# XII. APPENDIX B. DATA SET DESCRIPTION

The data set used in this study is a subsample of the 1973 Exact Match File, a nationally representative sample of all Americans in 1972. A respondent in the Match File was included in the subsample if she or he was a "good match," 62 or older in 1972, and received social security benefits in 1972. Two data sets were constructed: single and married.

The single data set included 353 respondents: 138 males (39 percent of all single respondents) and 215 females (61 percent of all single respondents). There are 2,771 couples included in the married data set, where at least one member of the couple satisfied the sorting criteria. The total number of respondents included in the study was 5,895. The following tables describe the characteristics of the data sets.

Table 12.1. Summary statistics

| | | |
|---|---|---|
| Total population | 5,895 | |
| Marital status | | |
| Married | 5,542 | (94% of sample) |
| Single | 353 | (6% of sample) |
| Race | | |
| White | 5,643 | (96% of sample) |
| Nonwhite | 252 | (4% of sample) |
| Men | | |
| Total | 2,909 | (49% of sample) |
| Marital status | | |
| Married | 2,771 | (95%) |
| Single | 138 | (5%) |
| Median age | | |
| Married | | |
| White | 69 | |
| Nonwhite | 69 | |
| Single | | |
| White | 69 | |
| Nonwhite | 69 | |
| Women | | |
| Total | 2,986 | (51% of sample) |
| Marital status | | |
| Married | 2,771 | (93%) |
| Single | 215 | (7%) |
| Median age | | |
| Married | | |
| White | 66 | |
| Nonwhite | 61 | |
| Single | | |
| White | 70 | |
| Nonwhite | 69 | |

Table 12.2. Age distribution by race, marital status, and sex

| Race, marital status, and sex | Age in 1972 | | | | | |
|---|---|---|---|---|---|---|
| | Less than 61 | 62-64 | 65 | 66-72 | More than 72 | Total |
| <u>White</u> | | | | | | |
| Married | | | | | | |
| Men | 50 | 282 | 197 | 1,344 | 783 | 2,656 |
| Women | 610 | 512 | 174 | 1,035 | 325 | 2,656 |
| Nonmarried | | | | | | |
| Men | 0 | 18 | 13 | 64 | 33 | 128 |
| Women | 0 | 13 | 12 | 125 | 53 | 203 |
| <u>Nonwhite</u> | | | | | | |
| Married | | | | | | |
| Men | 1 | 15 | 8 | 60 | 31 | 115 |
| Women | 60 | 12 | 4 | 32 | 7 | 115 |
| Nonmarried | | | | | | |
| Men | 0 | 2 | 0 | 7 | 1 | 10 |
| Women | 0 | 5 | 0 | 4 | 3 | 12 |
| | 721 | 859 | 408 | 2,671 | 1,236 | 5,985 |

Table 12.3. Year of retirement distribution by marital status, sex, and age

| Marital status, sex, and age | Year of retirement | | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | 1962 | 1963 | 1964 | 1965 | 1966 | 1967 | 1968 | 1969 | 1970 | 1971 | 1972 | 1973+ | Total |
| Nonmarried men | | | | | | | | | | | | | | |
| 62-64 | | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 3 | 11 | 6 | 0 | 20 |
| 65 | | 0 | 0 | 0 | 0 | 0 | 0 | 2 | 5 | 1 | 3 | 2 | 0 | 13 |
| 66-72 | | 1 | 1 | 10 | 7 | 7 | 14 | 7 | 9 | 7 | 4 | 4 | 0 | 71 |
| 73 and over | | 8 | 7 | 4 | 5 | 4 | 1 | 3 | 1 | 0 | 0 | 1 | 0 | 34 |
| | | 9 | 8 | 14 | 12 | 11 | 15 | 12 | 15 | 11 | 18 | 13 | 0 | 138 |
| Nonmarried women | | | | | | | | | | | | | | |
| 62-64 | | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 6 | 7 | 5 | 0 | 18 |
| 65 | | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 2 | 3 | 3 | 4 | 0 | 12 |
| 66-72 | | 5 | 8 | 7 | 9 | 19 | 15 | 24 | 13 | 14 | 10 | 5 | 0 | 129 |
| 73 and over | | 21 | 4 | 6 | 9 | 10 | 3 | 2 | 0 | 0 | 1 | 0 | 0 | 56 |
| | | 26 | 12 | 13 | 18 | 29 | 18 | 26 | 15 | 23 | 21 | 14 | 0 | 215 |
| Married men | | | | | | | | | | | | | | |
| 61 < | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 51 | 51 |
| 62-64 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 68 | 88 | 113 | 28 | 297 |
| 65 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 79 | 26 | 25 | 74 | 1 | 205 |
| 66-72 | 0 | 30 | 46 | 82 | 124 | 179 | 187 | 235 | 181 | 139 | 126 | 66 | 9 | 1,404 |
| 73 and over | 210 | 204 | 116 | 82 | 54 | 42 | 34 | 30 | 15 | 9 | 8 | 1 | 9 | 814 |
| | 210 | 234 | 162 | 164 | 178 | 221 | 221 | 265 | 275 | 242 | 247 | 254 | 98 | 2,771 |
| Married women | | | | | | | | | | | | | | |
| 61 < | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 670 | 670 |
| 62-64 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 119 | 161 | 159 | 85 | 524 |
| 65 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 112 | 32 | 14 | 18 | 2 | 178 |
| 66-72 | 0 | 54 | 74 | 110 | 131 | 158 | 158 | 175 | 85 | 57 | 34 | 15 | 16 | 1,067 |
| 73 and over | 120 | 69 | 49 | 26 | 20 | 17 | 2 | 6 | 7 | 4 | 5 | 1 | 6 | 332 |
| | 120 | 123 | 123 | 136 | 151 | 175 | 160 | 181 | 204 | 212 | 214 | 193 | 719 | 2,771 |

Table 12.4. Distribution by years of school completed and family income in 1972, men only

| Years of school completed | Family income in 1972: Married men: $0-4,000 | Married men: $4,001-6,000 | Married men: $6,001-10,000 | Married men: $10,001+ | Nonmarried men: $0-4,000 | Nonmarried men: $4,001-6,000 | Nonmarried men: $6,001-10,000 | Nonmarried men: $10,001+ |
|---|---|---|---|---|---|---|---|---|
| Total number | 724 | 697 | 711 | 639 | 77 | 17 | 18 | 26 |
| Total percent | 26 | 25 | 26 | 23 | 56 | 12 | 13 | 19 |
| Elementary | | | | | | | | |
| Less than 8 years | 313 | 182 | 140 | 74 | 33 | 1 | 4 | 6 |
| 8 years | 211 | 215 | 200 | 134 | 21 | 9 | 6 | 6 |
| High school | | | | | | | | |
| 1-3 years | 94 | 137 | 124 | 105 | 11 | 2 | 0 | 1 |
| 4 years | 63 | 106 | 149 | 155 | 7 | 4 | 3 | 7 |
| College | | | | | | | | |
| 1-3 years | 25 | 31 | 45 | 59 | 2 | 1 | 3 | 1 |
| 4 or more | 18 | 26 | 53 | 112 | 3 | 0 | 2 | 5 |

Table 12.5. Distribution by years of school completed and family income in 1972, women only

| Years of school completed | Family income in 1972 | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Married women | | | | Nonmarried women | | | |
| | $0-4,000 | $4,001-6,000 | $6,001-10,000 | $10,001+ | $0-4,000 | $4,001-6,000 | $6,001-10,000 | $10,001+ |
| Total number | 724 | 697 | 711 | 639 | 100 | 41 | 38 | 36 |
| Total percent | 26 | 25 | 26 | 23 | 46 | 19 | 18 | 17 |
| Elementary | | | | | | | | |
| Less than 8 years | 241 | 154 | 85 | 40 | 21 | 3 | 4 | 3 |
| 8 years | 202 | 177 | 144 | 98 | 14 | 7 | 4 | 2 |
| High school | | | | | | | | |
| 1-3 years | 127 | 118 | 149 | 108 | 11 | 2 | 6 | 4 |
| 4 years | 120 | 174 | 230 | 214 | 36 | 20 | 15 | 11 |
| College | | | | | | | | |
| 1-3 years | 23 | 54 | 61 | 86 | 4 | 3 | 3 | 4 |
| 4 or more | 11 | 20 | 42 | 93 | 14 | 6 | 6 | 12 |

Table 12.6. Distribution of claim status by sex, marital status, and age

| Sex, marital status, and age | Claim status: Primary worker | Dependent spouse | Not collecting |
|---|---|---|---|
| <u>Men</u>: | | | |
| Married | | | |
| 61 < | 0 | 0 | 51 |
| 62-64 | 268 | 0 | 29 |
| 65 | 204 | 0 | 1 |
| 66-72 | 1,389 | 2 | 13 |
| 73+ | 803 | 0 | 11 |
| Nonmarried | | | |
| 62-64 | 20 | 0 | 0 |
| 65 | 13 | 0 | 0 |
| 66-72 | 71 | 0 | 0 |
| 73+ | 34 | 0 | 0 |
| <u>Women</u>: | | | |
| Married | | | |
| 61 < | 0 | 0 | 670 |
| 62-64 | 234 | 218 | 72 |
| 65 | 96 | 80 | 2 |
| 66-72 | 521 | 531 | 15 |
| 73+ | 147 | 179 | 6 |
| Nonmarried | | | |
| 62-64 | 18 | 0 | 0 |
| 65 | 12 | 0 | 0 |
| 66-72 | 129 | 0 | 0 |
| 73+ | 56 | 0 | 0 |
| Total | 4,015 | 1,010 | 870 |

# XIII. APPENDIX C. TABLES

Table 13.1. Annual return rate on U.S. government bonds and stock market, 1937–1972[a]

| Period | Annual nominal return rate: U.S. government bonds (1) | Annual nominal return rate: Stock market (2) | Average of columns (1) and (2) (3) | Average annual real rate of return |
|---|---|---|---|---|
| 1937 | 2.74 | 4.38 | 3.56 | -0.04 |
| 1938 | 2.61 | -24.93 | -11.16 | -9.26 |
| 1939 | 2.41 | 8.99 | 5.7 | 7.1 |
| 1940 | 2.26 | -3.71 | -0.725 | -1.725 |
| 1941 | 2.05 | -5.28 | -1.615 | -6.615 |
| 1942 | 2.46 | -5.79 | -1.665 | -12.365 |
| 1943 | 2.47 | 33.14 | 17.805 | 11.705 |
| 1944 | 2.48 | 12.91 | 7.695 | 5.995 |
| 1945 | 2.37 | 23.72 | 13.045 | 10.745 |
| 1946 | 2.19 | 15.89 | 9.04 | 0.54 |
| 1947 | 2.25 | -6.73 | -2.24 | -16.64 |
| 1948 | 2.44 | 8.13 | 5.285 | -2.515 |
| 1949 | 2.31 | 8.58 | 5.445 | 6.445 |
| 1950 | 2.32 | 25.18 | 13.75 | 12.75 |
| 1951 | 2.57 | 25.52 | 14.045 | 6.145 |
| 1952 | 2.68 | 14.73 | 8.705 | 6.505 |
| 1953 | 2.94 | 6.42 | 4.68 | 3.88 |
| 1954 | 2.55 | 23.06 | 12.805 | 12.305 |
| 1955 | 2.84 | 35.08 | 18.96 | 19.36 |
| 1956 | 3.08 | 9.20 | 6.14 | 4.64 |
| 1957 | 3.47 | 8.38 | 5.925 | 2.325 |
| 1958 | 3.43 | 8.16 | 5.795 | 3.095 |
| 1959 | 4.07 | 24.09 | 14.08 | 13.28 |
| 1960 | 4.01 | .90 | 2.455 | 0.855 |
| 1961 | 3.90 | 20.18 | 12.04 | 11.04 |
| 1962 | 3.95 | -2.68 | 0.635 | -0.465 |
| 1963 | 4.00 | 14.54 | 9.27 | 8.07 |
| 1964 | 4.15 | 18.24 | 11.195 | 9.895 |
| 1965 | 4.21 | 11.13 | 7.67 | 5.97 |
| 1966 | 4.66 | .24 | 2.45 | -0.45 |
| 1967 | 4.85 | 10.93 | 7.89 | 4.99 |
| 1968 | 5.25 | 10.31 | 7.78 | 3.58 |
| 1969 | 6.10 | 2.52 | 4.31 | -1.09 |
| 1970 | 6.59 | 3.84 | 5.215 | -0.685 |
| 1971 | 5.74 | 23.88 | 14.81 | 10.51 |
| 1972 | 5.63 | 13.63 | 9.63 | 6.33 |

[a]U.S. Bureau of the Census (1960, 1975).

Table 13.2. Consumer price index, U.S. city average, all items, 1937-1972[a] (1967 = 100)

| Year | Consumer price index: all items | Inflation rate (percent) |
|---|---|---|
| 1937 | 43.0 | 3.6 |
| 1938 | 42.2 | -1.9 |
| 1939 | 41.6 | -1.4 |
| 1940 | 42.0 | 1.0 |
| 1941 | 44.1 | 5.0 |
| 1942 | 48.8 | 10.7 |
| 1943 | 51.9 | 6.1 |
| 1944 | 52.7 | 1.7 |
| 1945 | 53.9 | 2.3 |
| 1946 | 58.5 | 8.5 |
| 1947 | 66.9 | 14.4 |
| 1948 | 72.1 | 7.8 |
| 1949 | 71.4 | -1.0 |
| 1950 | 72.1 | 1.0 |
| 1951 | 77.8 | 7.9 |
| 1952 | 79.5 | 2.2 |
| 1953 | 80.1 | 0.8 |
| 1954 | 80.5 | 0.5 |
| 1955 | 80.2 | -0.4 |
| 1956 | 81.4 | 1.5 |
| 1957 | 84.3 | 3.6 |
| 1958 | 86.6 | 2.7 |
| 1959 | 87.3 | 0.8 |
| 1960 | 88.7 | 1.6 |
| 1961 | 89.6 | 1.0 |
| 1962 | 90.6 | 1.1 |
| 1963 | 91.7 | 1.2 |
| 1964 | 92.9 | 1.3 |
| 1965 | 94.5 | 1.7 |
| 1966 | 97.2 | 2.9 |
| 1967 | 100.0 | 2.9 |
| 1968 | 104.2 | 4.2 |
| 1969 | 109.8 | 5.4 |
| 1970 | 116.3 | 5.9 |
| 1971 | 121.3 | 4.3 |
| 1972 | 125.3 | 3.3 |

[a]U.S. President (1976).

Table 13.3. OAI contribution and tax base, 1937-1972[a]

| Year | Wage and salary tax rate, $t_y$ | Self-employment tax rate, $t_{ys}$ | Taxable maximum, $M_y$ |
|---|---|---|---|
| 1937 | .0111 | 0 | 3000. |
| 1938 | .0110 | 0 | 3000. |
| 1939 | .0110 | 0 | 3000. |
| 1940 | .0111 | 0 | 3000. |
| 1941 | .0135 | 0 | 3000. |
| 1942 | .0133 | 0 | 3000. |
| 1943 | .0128 | 0 | 3000. |
| 1944 | .0124 | 0 | 3000. |
| 1945 | .0125 | 0 | 3000. |
| 1946 | .0140 | 0 | 3000. |
| 1947 | .0147 | 0 | 3000. |
| 1948 | .0159 | 0 | 3000. |
| 1949 | .0155 | 0 | 3000. |
| 1950 | .0243 | 0 | 3000. |
| 1951 | .0245 | .0184 | 3600. |
| 1952 | .0244 | .0183 | 3600. |
| 1953 | .0249 | .0187 | 3600. |
| 1954 | .0336 | .0252 | 3600. |
| 1955 | .0342 | .0256 | 4200. |
| 1956 | .0346 | .0259 | 4200. |
| 1957 | .0352 | .0264 | 4200. |
| 1958 | .0354 | .0266 | 4200. |
| 1959 | .0397 | .0298 | 4800. |
| 1960 | .0485 | .0364 | 4800. |
| 1961 | .0484 | .0363 | 4800. |
| 1962 | .0509 | .0383 | 4800. |
| 1963 | .0599 | .0446 | 4800. |
| 1964 | .0600 | .0446 | 4800. |
| 1965 | .0597 | .0445 | 4800. |
| 1966 | .0613 | .0462 | 6600. |
| 1967 | .0626 | .0474 | 6600. |
| 1968 | .0584 | .0447 | 7800. |
| 1969 | .0656 | .0492 | 7800. |
| 1970 | .0644 | .0483 | 7800. |
| 1971 | .0717 | .0538 | 7800. |
| 1972 | .0719 | .0540 | 9000. |

[a]Freiden, Leimer, and Hoffman (1976).

Table 13.4. Sex-neutral survivor probabilities

| Age | | Age | |
|---|---|---|---|
| 40[a] | .997 | 71 | .96037 |
| 41 | .9967 | 72 | .95704 |
| 42 | .99638 | 73 | .95334 |
| 43 | .99603 | 74 | .94923 |
| 44 | .99565 | 75 | .94471 |
| 45 | .99524 | 76 | .93977 |
| 46 | .99479 | 77 | .93435 |
| 47 | .99427 | 78 | .92835 |
| 48 | .99367 | 79 | .92169 |
| 49 | .99300 | 80 | .91441 |
| 50 | .99226 | 81 | .90652 |
| 51 | .99148 | 82 | .89798 |
| 52 | .99071 | 83 | .88878 |
| 53 | .98995 | 84 | .87890 |
| 54 | .98918 | 85 | .87826 |
| 55 | .98839 | 86 | .85686 |
| 56 | .98751 | 87 | .84478 |
| 57 | .98648 | 88 | .83209 |
| 58 | .98527 | 89 | .81891 |
| 59 | .98389 | 90 | .80540 |
| 60 | .98239 | 91 | .79153 |
| 61 | .98083 | 92 | .77751 |
| 62[b] | .97918 | 93 | .76370 |
| 63 | .97748 | 94 | .75031 |
| 64 | .97569 | 95 | .73732 |
| 65 | .97378 | 96 | .72494 |
| 66 | .97372 | 97 | .71355 |
| 67 | .97124 | 98 | .70333 |
| 68 | .96873 | 99 | .69443 |
| 69 | .96614 | 100 | .68653 |
| 70 | .96338 | 101 | .67910 |

[a]For ages 40-61, National Center for Health Statistics, Table 1 (1964).

[b]For ages 62 and older, Bayo (1972).

Table 13.5. Age-sex-race specific survivor probabilities

| Age | White Men | White Women | Nonwhite Men | Nonwhite Women |
|---|---|---|---|---|
| 40[a] | .99668 | .99810 | .99251 | .99439 |
| 41 | .99632 | .99791 | .99186 | .99389 |
| 42 | .99591 | .99771 | .99125 | .99344 |
| 43 | .99546 | .99748 | .99069 | .99304 |
| 44 | .99496 | .99724 | .99016 | .99267 |
| 45 | .99442 | .99697 | .98962 | .99231 |
| 46 | .99383 | .99669 | .98899 | .99186 |
| 47 | .99314 | .99638 | .98817 | .99125 |
| 48 | .99234 | .99604 | .98708 | .99043 |
| 49 | .99144 | .99568 | .98578 | .98942 |
| 50 | .99045 | .99527 | .98435 | .98833 |
| 51 | .98942 | .99483 | .98290 | .98721 |
| 52 | .98838 | .99440 | .98146 | .98608 |
| 53 | .98736 | .99399 | .98006 | .98496 |
| 54 | .98632 | .99358 | .97869 | .98383 |
| 55 | .98525 | .99313 | .97727 | .98269 |
| 56 | .98407 | .99260 | .97573 | .98148 |
| 57 | .9827 | .99195 | .97411 | .98017 |
| 58 | .98109 | .99114 | .97238 | .97870 |
| 59 | .97926 | .99019 | .97053 | .97713 |
| 60 | .97729 | .98912 | .96863 | .97541 |
| 61 | .97524 | .98797 | .96665 | .97368 |
| 62[b] | .97466 | .9[illegible]10 | .97301 | .98529 |
| 63 | .97244 | .98806 | .97073 | .98376 |
| 64 | .97003 | .98689 | .96823 | .98203 |
| 65 | .96742 | .98557 | .96558 | .98018 |
| 66 | .96528 | .98464 | .96148 | .97737 |
| 67 | .96207 | .98259 | .95772 | .97464 |
| 68 | .95882 | .98052 | .95397 | .97192 |
| 69 | .95549 | .97836 | .95024 | .96921 |
| 70 | .95201 | .97606 | .94650 | .96645 |
| 71 | .94833 | .97354 | .94268 | .96356 |
| 72 | .94440 | .97073 | .93870 | .96045 |
| 73 | .94018 | .96757 | .93451 | .95708 |
| 74 | .93560 | .96401 | .93004 | .95342 |
| 75 | .93061 | .96001 | .92527 | .94943 |
| 76 | .92518 | .95552 | .92022 | .94511 |
| 77 | .91929 | .95049 | .91488 | .94049 |

[a]For ages 40 to 61, National Center for Health Statistics, Tables 5-9 (1964).

[b]For ages 62 and older, Bayo (1972).

Table 13.5. continued

| Age | White Men | White Women | Nonwhite Men | Nonwhite Women |
|---|---|---|---|---|
| 78 | .91289 | .94485 | .90925 | .93557 |
| 79 | .90593 | .93855 | .90326 | .93042 |
| 80 | .89836 | .93154 | .89690 | .92504 |
| 81 | .89018 | .92381 | .89015 | .91941 |
| 82 | .88131 | .91532 | .88309 | .91353 |
| 83 | .87172 | .90605 | .87574 | .90733 |
| 84 | .86144 | .89597 | .86803 | .90078 |
| 85 | .85043 | .88507 | .86000 | .89376 |
| 86 | .83865 | .87334 | .85183 | .88628 |
| 87 | .82610 | .86086 | .84357 | .87843 |
| 88 | .81276 | .84774 | .83502 | .87026 |
| 89 | .79861 | .83414 | .82626 | .86181 |
| 90 | .78389 | .82005 | .81773 | .85341 |
| 91 | .76880 | .80551 | .80995 | .84531 |
| 92 | .75358 | .79050 | .80329 | .83785 |
| 93 | .73865 | .77522 | .79782 | .83118 |
| 94 | .72419 | .75987 | .79349 | .82519 |
| 95 | .71002 | .73031 | .79020 | .81929 |
| 96 | .69570 | .71678 | .78780 | .81277 |
| 97 | .68189 | .70438 | .78578 | .80543 |
| 98 | .66897 | .69313 | .78296 | .79758 |
| 99 | .65743 | .68316 | .77940 | .78979 |
| 100 | .64690 | .67395 | .77496 | .78227 |
| 101 | .63610 | .67395 | .76960 | .77523 |

# XIV. APPENDIX D. COMPARISON OF COMPOUNDING SCHEMES

Table 14.1. Comparison of accumulated contributions

| Family income level in 1972 | Mean difference between ROATC and TATC[a] | Percentage of population |
|---|---|---|
| 0 - 1,000 | $-276.00 | .9 |
| 1,001 - 1,500 | -134.00 | 1.1 |
| 1,501 - 2,000 | -185.00 | 2.7 |
| 2,001 - 2,500 | -206.00 | 3.9 |
| 2,501 - 3,000 | -233.00 | 4.9 |
| 3,001 - 3,500 | -216.00 | 5.6 |
| 3,501 - 4,000 | -252.00 | 7.4 |
| 4,001 - 5,000 | -306.00 | 14.2 |
| 5,001 - 6,000 | -349.00 | 12.2 |
| 6,001 - 8,000 | -391.00 | 16.0 |
| 8,001 - 10,000 | -467.00 | 9.4 |
| 10,001 - 20,000 | -501.00 | 16.5 |
| 20,001+ | -489.00 | 5.2 |
| Total | $-308.00 | 100 |

[a]ROATC is the beneficiary's accumulated contributions credited to his/her account using the roll-over compounding scheme. TATC is the beneficiary's accumulated contributions based on the traditional compounding scheme.

# XV. APPENDIX E. RESULTS

Table 15.1. Aggregate data for Table 7.1

| | Total OAI benefits in 1972 | Type-6 actuarially fair benefits (total) | Total amount of intergenerational transfers | Percent of[a] total intergenerational transfers | Cumulative[a] percent |
|---|---|---|---|---|---|
| 0- 1,000 | 6,975 | 174 | 6,801 | 0.11 | 0.11 |
| 1,001- 1,500 | 43,673 | 3,127 | 40,546 | 0.64 | 0.75 |
| 1,501- 2,000 | 104,007 | 9,010 | 94,997 | 1.51 | 2.26 |
| 2,001- 2,500 | 236,281 | 20,545 | 215,736 | 3.43 | 5.69 |
| 2,501- 3,000 | 302,978 | 28,388 | 274,590 | 4.63 | 10.32 |
| 3,001- 3,500 | 387,296 | 41,054 | 346,242 | 5.50 | 15.82 |
| 3,501- 4,000 | 575,508 | 65,193 | 510,315 | 8.11 | 23.93 |
| 4,001- 5,000 | 1,044,646 | 120,014 | 924,632 | 14.69 | 38.62 |
| 5,001- 6,000 | 887,077 | 107,522 | 779,555 | 12.39 | 51.01 |
| 6,001- 8,000 | 1,190,461 | 147,815 | 1,042,646 | 16.57 | 67.58 |
| 8,001-10,000 | 702,368 | 82,856 | 619,512 | 9.84 | 77.42 |
| 10,001-20,000 | 1,242,319 | 131,220 | 1,111,099 | 17.66 | 95.08 |
| 20,001+ | 366,117 | 39,503 | 326,614 | 5.19 | 100.27 |
| Total | 7,089,706 | 796,421 | 6,293,285 | 100.27 | |

[a]Totals may not add to 100 because of rounding.

Table 15.2. Changes in the percentage of redistribution due to indexing for married, both retired households

| Total family income in 1972 | Gender-merged, earnings adjusted | | | Sex-race-distinct, earnings adjusted | | |
|---|---|---|---|---|---|---|
| | Type-1 | Type-4 | (2)-(1) | Type-2 | Type-5 | (4)-(3) |
| | (1) Nonindexed[a] | (2) Indexed[a] | Change in percentage of redistribution | (3) Nonindexed[a] | (4) Indexed[a] | Change in percentage of redistribution |
| 0- 1,000 | 97.4 | 97.7 | 0.3 | 97.5 | 97.8 | 0.3 |
| 1,001- 1,500 | 93.1 | 93.9 | 0.8 | 93.3 | 94.1 | 0.8 |
| 1,501- 2,000 | 90.7 | 91.8 | 1.1 | 90.9 | 92.0 | 1.1 |
| 2,001- 2,500 | 90.4 | 91.3 | 0.9 | 90.5 | 91.6 | 1.1 |
| 2,501- 3,000 | 89.6 | 90.6 | 1.0 | 89.7 | 91.0 | 1.3 |
| 3,001- 3,500 | 88.2 | 89.4 | 1.2 | 88.4 | 89.7 | 1.3 |
| 3,501- 4,000 | 87.4 | 88.6 | 1.2 | 87.6 | 88.9 | 1.3 |
| 4,001- 5,000 | 86.8 | 88.2 | 1.4 | 87.1 | 88.6 | 1.5 |
| 5,001- 6,000 | 86.0 | 87.5 | 1.5 | 86.3 | 87.9 | 1.6 |
| 6,001- 8,000 | 85.4 | 87.1 | 1.7 | 85.7 | 87.5 | 1.8 |
| 8,001-10,000 | 85.9 | 87.7 | 1.8 | 86.2 | 88.1 | 1.9 |
| 10,001-20,000 | 87.4 | 88.9 | 1.5 | 87.6 | 89.3 | 1.7 |
| 20,001+ | 87.1 | 88.6 | 1.5 | 87.3 | 88.9 | 1.6 |

[a]Raw data used to calculate the percentage of redistribution for each family income classification is available upon request.

Table 15.3. Changes in the percentage of redistribution under different survivorship probability assumptions, nonearning test adjusted for married, both retired households

| Total family income in 1972 | Annuity-type, indexed | | | Change in percentage of redistribution | |
|---|---|---|---|---|---|
| | Type-1[a] (4) | Type-2[a] (5) | Type-3[a] (6) | (5)-(4) | (6)-(4) |
| 0- 1,000 | 97.7 | 97.8 | 97.6 | 0.1 | -0.1 |
| 1,001- 1,500 | 93.9 | 94.1 | 92.9 | 0.2 | -1.0 |
| 1,501- 2,000 | 91.8 | 92.0 | 91.3 | 0.2 | -0.5 |
| 2,001- 2,500 | 91.1 | 91.4 | 91.1 | 0.3 | 0.0 |
| 2,501- 3,000 | 90.3 | 90.6 | 90.3 | 0.3 | 0.0 |
| 3,001- 3,500 | 89.1 | 89.5 | 89.1 | 0.4 | 0.0 |
| 3,501- 4,000 | 88.3 | 88.6 | 88.4 | 0.3 | 0.1 |
| 4,001- 5,000 | 87.9 | 88.3 | 88.2 | 0.4 | 0.3 |
| 5,001- 6,000 | 87.1 | 87.6 | 87.5 | 0.5 | 0.4 |
| 6,001- 8,000 | 86.4 | 86.8 | 86.8 | 0.4 | 0.4 |
| 8,001-10,000 | 86.6 | 87.1 | 87.1 | 0.5 | 0.5 |
| 10,001-20,000 | 87.1 | 87.5 | 87.6 | 0.4 | 0.5 |
| 20,001+ | 87.0 | 87.4 | 87.7 | 0.4 | 0.7 |
| Mean | 87.6 | 88.0 | 88.0 | 0.4 | 0.4 |

[a]Raw data used to calculate the percentage of redistribution for each family income classification is available upon request.

Table 15.4. Summary percentage point comparisons for married, both retired households by annuity type, sex, and household type

| | Type-1 | | Type-2 | | Type-3 | | Type-4 | | Type-5 | | Type-6 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Two earner | One earner | Two earner | One earner | Two earner | One earner | Two earner | One earner | Two earner | One earner | Two earner | One earner |
| Poorest to richest percentage point gap | | | | | | | | | | | | |
| Female | 6 | 7 | 6 | 6 | 6 | 6 | 6 | 6 | 5 | 7 | 4 | 6 |
| Male | 7 | 3 | 6 | 3 | 6 | 3 | 6 | 3 | 6 | 4 | 5 | 3 |
| Household unit | 7 | 4 | 6 | 4 | 6 | 4 | 6 | 4 | 6 | 4 | 5 | 4 |
| Highest percentage of redistribution | | | | | | | | | | | | |
| Female | 87 | 90 | 88 | 90 | 88 | 90 | 89 | 91 | 89 | 92 | 89 | 92 |
| Male | 93 | 96 | 93 | 96 | 93 | 96 | 94 | 96 | 94 | 97 | 94 | 97 |
| Household unit | 91 | 94 | 91 | 94 | 91 | 94 | 92 | 95 | 92 | 95 | 92 | 95 |
| Lowest percentage of redistribution | | | | | | | | | | | | |
| Female | 78 | 79 | 79 | 79 | 79 | 79 | 81 | 81 | 81 | 81 | 81 | 81 |
| Male | 86 | 91 | 86 | 91 | 86 | 91 | 87 | 92 | 88 | 92 | 88 | 92 |
| Household unit | 83 | 87 | 83 | 87 | 84 | 88 | 85 | 88 | 85 | 89 | 86 | 89 |

Table 15.5. Male to female differences in percentage of redistribution controlling for family income and family type

| Total family income in 1972 | Type-1 | | | Type-2 | | | Type-3 | | |
|---|---|---|---|---|---|---|---|---|---|
| | Female | Male | Difference | Female | Male | Difference | Female | Male | Difference |
| <u>Two earner</u> | | | | | | | | | |
| 0- 2,000 | 87 | 93 | +6 | 88 | 93 | +5 | 88 | 93 | +5 |
| 2,001- 2,500 | 88 | 93 | +5 | 88 | 93 | +5 | 88 | 93 | +5 |
| 2,501- 3,000 | 88 | 92 | +4 | 88 | 92 | +4 | 88 | 92 | +4 |
| 3,001- 3,500 | 84 | 88 | +4 | 85 | 88 | +3 | 84 | 88 | +4 |
| 3,501- 4,000 | 81 | 88 | +7 | 82 | 89 | +7 | 82 | 89 | +7 |
| 4,001- 5,000 | 80 | 87 | +7 | 80 | 87 | +7 | 80 | 87 | +7 |
| 5,001- 6,000 | 78 | 87 | +9 | 79 | 87 | +8 | 79 | 87 | +8 |
| 6,001- 8,000 | 79 | 86 | +7 | 80 | 86 | +6 | 80 | 86 | +6 |
| 8,001-10,000 | 80 | 86 | +6 | 81 | 86 | +5 | 81 | 86 | +5 |
| 10,001-20,000 | 84 | 87 | +3 | 84 | 87 | +3 | 84 | 88 | +4 |
| 20,001+ | 81 | 86 | +5 | 82 | 87 | +5 | 82 | 87 | +5 |
| Mean | | | 5.7 | | | 5.3 | | | 5.5 |
| <u>One earner</u> | | | | | | | | | |
| 0- 2,000 | 90 | 96 | +6 | 90 | 96 | +6 | 90 | 96 | +6 |
| 2,001- 2,500 | 90 | 96 | +6 | 90 | 96 | +6 | 91 | 96 | +5 |
| 2,501- 3,000 | 85 | 94 | +9 | 86 | 94 | +8 | 86 | 94 | +8 |
| 3,001- 3,500 | 84 | 93 | +9 | 85 | 93 | +8 | 85 | 93 | +8 |
| 3,501- 4,000 | 82 | 92 | +10 | 82 | 92 | +10 | 82 | 92 | +10 |
| 4,001- 5,000 | 83 | 92 | +9 | 83 | 93 | +10 | 83 | 93 | +10 |
| 5,001- 6,000 | 79 | 91 | +12 | 79 | 91 | +12 | 79 | 91 | +12 |
| 6,001- 8,000 | 80 | 91 | +11 | 81 | 91 | +10 | 81 | 91 | +10 |
| 8,001-10,000 | 79 | 91 | +12 | 79 | 91 | +12 | 79 | 91 | +12 |
| 10,001-20,000 | 82 | 92 | +10 | 82 | 92 | +10 | 83 | 92 | +9 |
| 20,001+ | 83 | 93 | +10 | 84 | 93 | +9 | 84 | 93 | +9 |
| Mean | | | 9.6 | | | 9.2 | | | 9.0 |

Table 15.6. Family type differences in percentage of redistribution controlling for family income and sex

| Total family income in 1972 | Type-1 | | | Type-2 | | | Type-3 | | |
|---|---|---|---|---|---|---|---|---|---|
| | Two earner | One earner | Dif-ference | Two earner | One earner | Dif-ference | Two earner | One earner | Dif-ference |
| **Females** | | | | | | | | | |
| 0- 2,000 | 87 | 90 | +3 | 88 | 90 | +2 | 88 | 90 | +2 |
| 2,001- 2,500 | 88 | 90 | +2 | 88 | 90 | +2 | 88 | 91 | +3 |
| 2,501- 3,000 | 88 | 85 | -3 | 88 | 86 | -2 | 88 | 86 | -2 |
| 3,001- 3,500 | 84 | 84 | 0 | 85 | 85 | 0 | 84 | 85 | +1 |
| 3,501- 4,000 | 81 | 82 | +1 | 82 | 82 | 0 | 82 | 82 | 0 |
| 4,001- 5,000 | 80 | 83 | +3 | 80 | 83 | +3 | 80 | 83 | +3 |
| 5,001- 6,000 | 78 | 79 | +1 | 79 | 79 | 0 | 79 | 79 | 0 |
| 6,001- 8,000 | 79 | 80 | +1 | 80 | 81 | +1 | 80 | 81 | +1 |
| 8,001-10,000 | 80 | 79 | -1 | 81 | 79 | -2 | 81 | 79 | -2 |
| 10,001-20,000 | 84 | 82 | -2 | 84 | 82 | -2 | 84 | 83 | -1 |
| 20,001+ | 81 | 83 | +2 | 82 | 84 | +2 | 82 | 84 | +2 |
| Mean | | | | | | | | | |
| **Males** | | | | | | | | | |
| 0- 2,000 | 93 | 96 | +3 | 93 | 96 | +3 | 93 | 96 | +3 |
| 2,001- 2,500 | 93 | 96 | +3 | 93 | 96 | +3 | 93 | 96 | +3 |
| 2,501- 3,000 | 92 | 94 | +2 | 92 | 94 | +2 | 92 | 94 | +2 |
| 3,001- 3,500 | 88 | 93 | +5 | 88 | 93 | +5 | 88 | 93 | +5 |
| 3,501- 4,000 | 88 | 92 | +4 | 89 | 92 | +3 | 89 | 92 | +3 |
| 4,001- 5,000 | 87 | 92 | +5 | 87 | 93 | +6 | 87 | 93 | +6 |
| 5,001- 6,000 | 87 | 91 | +4 | 87 | 91 | +4 | 87 | 91 | +4 |
| 6,001- 8,000 | 86 | 91 | +5 | 86 | 91 | +5 | 86 | 91 | +5 |
| 8,001-10,000 | 86 | 91 | +5 | 86 | 91 | +5 | 86 | 91 | +5 |
| 10,001-20,000 | 87 | 92 | +5 | 87 | 92 | +5 | 88 | 92 | +4 |
| 20,001+ | 86 | 93 | +7 | 87 | 93 | +6 | 87 | 93 | +6 |
| Mean | | | | | | | | | |

Table 15.7. Nonindexed to indexed differences in percentage of redistribution controlling for family income and household unit

| | Type-1 | | | Type-2 | | | Type-3 | | |
|---|---|---|---|---|---|---|---|---|---|
| | Non-indexed | Indexed | Dif-ferences | Non-indexed | Indexed | Dif-ferences | Non-indexed | Indexed | Dif-ferences |
| <u>Two-earner</u> | | | | | | | | | |
| 0- 2,000 | 91 | 92 | +1 | 91 | 92 | +1 | 91 | 92 | +1 |
| 2,001- 2,500 | 91 | 92 | +1 | 92 | 93 | +1 | 92 | 93 | +1 |
| 2,501- 3,000 | 90 | 91 | +1 | 90 | 91 | +1 | 90 | 91 | +1 |
| 3,001- 3,500 | 87 | 88 | +1 | 87 | 88 | +1 | 87 | 88 | +1 |
| 3,501- 4,000 | 86 | 87 | +1 | 86 | 87 | +1 | 86 | 87 | +1 |
| 4,001- 5,000 | 84 | 86 | +2 | 85 | 86 | +1 | 85 | 86 | +1 |
| 5,001- 6,000 | 84 | 85 | +1 | 84 | 86 | +2 | 84 | 86 | +2 |
| 6,001- 8,000 | 83 | 85 | +2 | 83 | 85 | +2 | 84 | 86 | +2 |
| 8,001-10,000 | 83 | 85 | +2 | 84 | 86 | +2 | 84 | 86 | +2 |
| 10,001-20,000 | 86 | 87 | +1 | 86 | 87 | +1 | 86 | 88 | +2 |
| 20,001+ | 84 | 86 | +2 | 85 | 86 | +1 | 85 | 87 | +2 |
| <u>One-earner</u> | | | | | | | | | |
| 0- 2,000 | 94 | 95 | +1 | 94 | 95 | +1 | 94 | 95 | +1 |
| 2,001- 2,500 | 94 | 95 | +1 | 94 | 95 | +1 | 95 | 95 | 0 |
| 2,501- 3,000 | 91 | 92 | +1 | 91 | 92 | +1 | 91 | 92 | +1 |
| 3,001- 3,500 | 91 | 91 | 0 | 91 | 92 | +1 | 91 | 92 | +1 |
| 3,501- 4,000 | 89 | 90 | +1 | 89 | 90 | +1 | 89 | 90 | +1 |
| 4,001- 5,000 | 90 | 90 | 0 | 90 | 91 | +1 | 90 | 91 | +1 |
| 5,001- 6,000 | 87 | 88 | +1 | 87 | 89 | +2 | 87 | 89 | +2 |
| 6,001- 8,000 | 88 | 89 | +1 | 89 | 89 | 0 | 88 | 89 | +1 |
| 8,001-10,000 | 87 | 89 | +2 | 88 | 89 | +1 | 88 | 89 | +1 |
| 10,001-20,000 | 89 | 90 | +1 | 89 | 90 | +1 | 89 | 90 | +1 |
| 20,001+ | 90 | 91 | +1 | 90 | 91 | +1 | 90 | 91 | +1 |

# XVI. APPENDIX F. DISAGGREGATION OF THE 1937 to 1950 REPORTED EARNINGS MEASURE

To correctly calculate the lifetime earnings measure, the 1937-1950 summary taxable earnings measure had to be disaggregated into year-specific reported earnings measures. This was accomplished by using the year-specific estimated annual quarters of coverage from 1937 to 1950 and the 1937-1950 summary taxable earnings measure. The following procedure was employed to estimate the year-specific reported earnings for 1937 to 1950. First, the estimated reported earnings for year i ($EREP_i$) was calculated by

$$EREP_i = \left(\frac{EQC_i}{TEQC}\right) \left(\frac{w_i}{\sum_{i=1937}^{50} \frac{w_i}{14}}\right) (TOTAL50) \tag{16.1}$$

where $EQC_i$ equals the estimated quarter of coverage in year i, TEQC equals the total estimated quarters of coverage for 1937 to 1950, $w_i$ equals the average annual earnings for full-time employee in manufacturing in year i, $\sum_{i=1937}^{50} w_i/14$ equals the average annual earnings for full-time employee in manufacturing over the 1937-1950 time period, and TOTAL50 equals the total reported earnings for the 1937-1950 time period, as reported on the Longitudinal Exact Match File. Hence, the estimated reported earnings are divided over the time interval proportionally to the estimated annual quarters of coverage and average annual earnings in manufacturing from 1937 to 1950.

Because the estimated reported earnings measures were adjusted for the changes in average earnings over time, the sum of the estimated reported earnings measures will not, in all likelihood, equal the total reported earnings reported in the Longitudinal Exact Match File. The estimation error is

$$BIAS = TOTAL50 - \sum_{i=1937}^{50} EREP. \tag{16.2}$$

The estimation error may be positive or negative depending on the location of the estimated quarters of coverage over the 1937-1950 time interval. The worker's estimated reported earnings are proportionally adjusted by the estimation error. That is, the estimation error is spread over the time period so as to preserve the proportion of estimated reported earnings in year i to the total estimated reported earnings from 1937 to 1950. The proportion of estimated reported earnings in year i ($EREP_i$) to the total estimated reported earnings from 1937 to 1950 is represented by

$$PRO_i = \frac{EREP_i}{\sum_{i=1937}^{1950} EREP_i} \tag{16.3}$$

for i equal to 1937 to 1950. The adjustment factor for each year ($ADJ_i$) is

$$ADJ_i = PRO_i \times BIAS \tag{16.4}$$

for i equal to 1937 to 1950. Finally, the adjustment factor for each year is used to adjust the estimated reported earnings for the same year. Hence, the reported earnings for year i ($REP_i$) is

$$REP_i = EREP_i + ADJ. \tag{16.5}$$