A General Model of Labor-Market Behavior of Older Persons

by Marjorie Honig and Giora Hanoch*

Identifying the separate effects of age, time period, and birth cohort is of obvious importance in studies of the older population and especially critical in analyzing the labor-market behavior of the elderly. Age and the aging process in particular are fundamentally associated with changes occurring in earnings and incomes of persons at or near retirement. This article estimates and analyzes the separate effects of age, cohort, and period for some key labor-market variables, utilizing panel data from the Retirement History Study of the Social Security Administration, matched with social security earnings records. The analysis applies a strategy recently developed by the authors for estimating these separate effects from period-cohort means. The approach provides a solution to the well-known identification problem that has long been an obstacle in these studies. Integrated with the age-period-cohort analysis is a behavioral simultaneous model of labor-force participation, annual hours and weeks of work, threshold labor-supply quantities and reservation wages, wage offers, and asset holdings of older married men and women who are heads of households.

Several factors distinguish the labor-supply decisions of older persons from those of the prime-age labor force: The availability of income maintenance in the form of old-age benefits from the social security program and other public support programs, income from other private and government pension plans and from their own resources accumulated for old age, compulsory retirement provisions, changing preferences for leisure, deteriorating health, and constraints on the demand side regarding the availability of part-time employment in semiretirement.

The standard theoretical and empirical models of the labor market developed for the analysis of prime-age workers must be adapted to these special features relevant to the older population. The social security old-age benefits program, for example, entails an earnings test that imposes an implicit tax on earnings (in addition to explicit income taxes). This feature gives rise to a segmented, noncontinuous individual labor-supply curve. The problems of selectivity and endogeneity associated with the individual's choice of budget segment under these conditions require an appropriate econometric methodology. Other provisions of the social security program increase the complexity of the analysis—for example, past earnings determine current eligibility and current earnings affect future social security benefits. The existence of various other pension plans and work-related sources of income further complicate matters. Market wages, nonwage income, and asset holdings of individuals are correlated in the cross-section with permanent home productivity and preferences for leisure and therefore with current labor-supply behavior.

The treatment of these variables as endogenous and the use of some additional simplifying assumptions make it possible to apply the static one-period labor-supply model to the analysis of older persons. Life-cycle and dynamic effects are particularly important for this segment of the population.

This article describes a longitudinal analysis of a simultaneous model of labor supply, market-wage determination, and the asset-accumulation behavior of the older population that involves estimates of the factors determining labor-force participation, annual hours and weeks of work, threshold labor-supply quantities and reservation wages, wage offers, and asset holdings. The analysis

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combines cross-section and overtime data using two merged microdata files (white married men and white women who are heads of families), each including several hundred variables created from interview waves (1969, 1971, and 1973) of the Retirement History Study (RHS) of the Social Security Administration. This study is the largest and most detailed survey on the older population that is currently available. The survey data are merged here with the earnings history records of the Social Security Administration.

The empirical methodology focuses on a longitudinal analysis that combines cross-section and time-series data. The analysis of data of this type has occupied an increasingly central role in a variety of social and economic studies. A focal problem in many of these studies has been the estimation of the separate effects of time period, cohort of birth or vintage, and age. The present discussion offers a solution to the well-known identification problem that arises because current year, year of birth, and age are linearly dependent. Empirical results are presented for the separate effects of age, cohort, and period in selected labor-market variables for a sample of older married men.

The estimated model, incorporating the analysis of age, cohort, and period, is designed to examine current patterns of labor supply and asset-accumulation behavior within the older white population in the United States. The model can also be used to predict (1) the effects of various proposed changes in social policy regarding the aged, such as modifications in the benefit structure and earings test in the social security program, (2) the impact of expanded coverage and benefits under private pension programs, (3) the effect of the recent increase in the compulsory retirement age, and (4) the effect of recent efforts to expand the part-time job market and to reduce the constraints on job mobility for older persons.

Labor-Market Behavior of Older Persons: A Theoretical Framework

The model discussed briefly in this section is based on a theoretical analysis of the individual labor-supply curve under conditions of an income-maintenance program such as social security.[1] The simultaneous model consists of the following equations—the probability of participation, annual hours of work, annual weeks, market wage, reservation wage, and asset holdings in each of the 3 years—allowing for variations in parameters over time. The variance-covariance matrix of residuals is partitioned into the permanent components that account for constant differences among individuals in preferences and in productivity, and the transitory components that are regarded as random.

The methodology used for the estimation is chosen with the aim of obtaining consistent but inexpensive estimates, sacrificing some efficiency relative to infeasible or very expensive maximum likelihood methods. The simpler estimation methods may be more robust, however, when the models are imperfectly specified. These simpler methods allow more intensive specification search at a lower cost in order to choose the best proxies for given conceptual variables from sets of alternative variables and better exclusion restrictions and functional forms. This specification search may yield higher returns than high-powered and costly estimation methods applied to restrictive predetermined specifications of the equations and definitions of the variables. A byproduct will be added insight about the unique data set developed for this study, which may be useful in a wide variety of analyses concerning the older population. Recommendations for improvements in future microsurveys may also be derived.

This large microdata file includes several hundred variables created from the 1969, 1971, and 1973 interview waves of the RHS. The file contains several alternative measures for each of many important variables such as wages, labor supply, labor-force experience (including job tenure and general market experience), types of assets and incomes, health, retirement status, pension and social security eligibility status, household and family characteristics, and individual background characteristics, as well as measures of changes over a period of time in these variables. The data were carefully screened for reporting and recording errors that cause extreme outlying observations, and considerable effort was devoted to avoiding exclusions from the sample based on missing information. All variables have been transformed from coded to quantitative form (including many dummy variables). The processed data are stored separately in two merged files, for white married men and white women who are heads of families (aged 58-63 in 1969)—3,683 and 1,514 observations, respectively.

The survey data have been merged with Social Security Administration record data on annual earnings and quarters of coverage of respondents and spouses for 1951-74. These data provide information on the work histories of respondents from age 40, including movements in and out of the labor force (specifically, social-security-covered employment), as well as checks for accuracy in the survey data on earnings and employment. In addition, it is possible to estimate with a high degree of accuracy the potential current and future retirement benefits available to respondents and spouses under the social security program.

The theoretical development in Hanoch and Honig[2] proposes a simple model with one individual, one period, two goods (leisure and other consumption), and exogenous wage rate, nonwage income, and benefits. Extensions

[1] See Giora Hanoch and Marjorie Honig, "The Labor-Supply Curve Under Income Maintenance Programs," **Journal of Public Economics,** February 1978.

[2] **Ibid.**

of the model account for the two-tired tax system in the social security program in effect in the years covered by the first three RHS panels, as well as for other complexities in social security provisions such as the effects of current earnings and benefits on future benefits, the endogeneity of wage rates and nonwage incomes, and intrafamily substitution effects. These effects are incorporated into the methodology developed for the empirical application of the model.

A brief description of the main findings of the model follows.[3] The supply function of hours H^* (W, Y) is uniquely associated with the utility function $U(X, L)$, where L is leisure $(L \leqslant T)$ and X is goods, W is a fixed hourly wage, and Y a given nonwage income.

Given a benefit B, an earnings maximum M, and an implicit tax rate t on benefits $(0 \leqslant t \leqslant 1)$ for earnings above M, the supply curve can take two alternative forms:

(1) If $t \geqslant t_0$, where t_0 is an endogenous threshold critical tax rate, the supply curve is of the general form shown in chart 1.

(2) If $t < t_0$, the supply curve is as shown in chart 2.

The backward-bending segment $H = \frac{M}{W}$ corresponds to the wage range where the optimal solution is at the corner. The individual works just enough to earn the maximum exempt amount $M = HW$.

[3] For proofs and details, see Giora Hanoch and Marjorie Honig, **ibid.**, and, for an illustration related to the Israeli social security system, see Giora Hanoch and Marjorie Honig, **The Effect of Social Security on Labor Supply** (Working Paper 88), Hebrew University, Department of Economics, Jerusalem, 1976.

Chart 1.—The individual labor supply curve, $t \geqslant t_0$

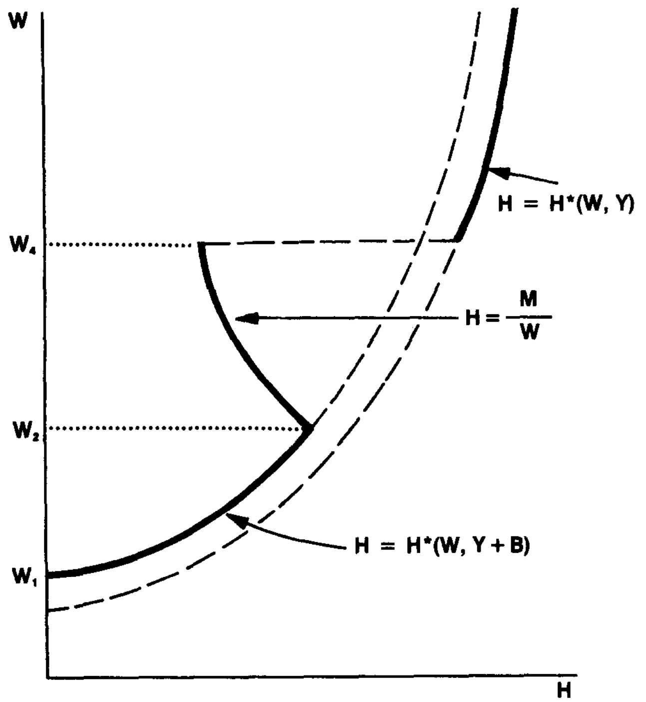
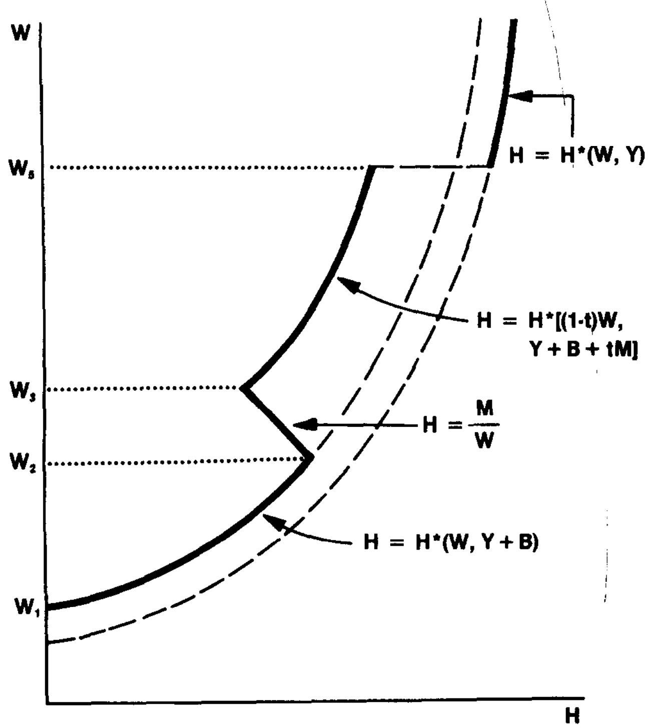


Chart 2.—The individual labor supply curve, $t < t_0$



The discontinuity (at W_4 or W_5) occurs where the worker is just indifferent between choosing the receipt of benefits or working and earning more and thereby opting out of the benefit system. The horizontal distance between H^* (W, Y) and this supply curve represents the total effect of the program where the difference H^* $(W, Y) - H^*$ $(W, Y + B)$ is the pure benefit effect, and the remaining difference (which may be negative) is the effect of the earnings test and the implicit tax. Changes in t above t_0 do not have any effect, and changes below t_0 may increase or decrease work, depending on the individual's location on the wage axis (relative to the supply curve, determined jointly by his preferences and the program parameters).

If the fixed money costs C are involved in entering the labor market, the supply curve will have a discontinuity at the point of entry, and the lowest positive segment of the supply curve will be:

$$H = H^*(W, Y-C+B) \geqslant H_0 \text{ (for } W > W_1^* > W_1)$$

where the minimum supply H_0 equates the utility $U(Y+B, T)$ when retired, with $U(Y-C+B+W_1{}^*H_0, T-H_0)$, when working H_0 hours, receiving a wage equal to the reservation wage $W_1{}^*$.

In the empirical application, individual variation in preferences is captured by introducing the effects of various measured variables Z into the basic supply function H^*, and by an additive stochastic component ϵ; that is,

$$H^* = H^*(W, Y; Z) + \epsilon \cdot$$

Z and ϵ introduce variation among individuals both in the location and form of each segment of the supply curve as well as in the critical wage rates that determine its relevant boundaries. Specification of H^* as additive in log w, y, z and ϵ facilitates the estimation of the function in all the segments jointly, provided the choice of the relevant segment is accounted for as an endogenous choice.

The estimation by maximum likelihood of the complete simultaneous model, including the probabilities of location on each segment and of retirement or nonparticipation, the hours (and perhaps weeks) equation, and the reservation wage as well as the wage equation, is clearly infeasible. Approximate two-stage limited information estimations may be carried out but are also costly and complex.

The considerably cheaper but consistent route to be taken in this study for estimating the model may be sketched as follows:

(1) Estimate the probability of participation by maximum likelihood probit analysis, disregarding at this stage the location of workers on other segments of the supply function and the quantity of labor.

(2) Estimate conditional probabilities among workers for locating on different segments of the supply function by a multinomial probit (or logit) estimation—or by a series of dichotomous probit estimates (if this yields reasonable results—such that probabilities are positive fractions adding to one for almost all individuals).

(3) Use the inverse Mill's ratios obtained from the probit estimates as additional explanatory variables in each behavioral equation as used in labor-supply estimates by Heckman and Hanoch.[4] Introducing these estimates variables into the equations corrects for selectivity biases implied by the endogenous choice of segment and their coefficients give estimates for some parameters of the residual variance-covariance matrix.

(4) The function H^* can be estimated by pooling all the noncorner solution cases together, if the appropriate wage and income variables are used as arguments (for individuals whose benefits are reduced via the earnings test at the $t = 0.5$ segment, for example, the relevant wage is $0.5W$, and the relevant income is $(Y + B + 0.5M)$) and if the relevant bias-correcting factors (for truncation at both ends of an interior segment) are added to the list of right-hand side variables.

(5) The linearized equations, corrected for all the relevant selection biases, are now estimated by linear methods applicable to a linear simultaneous model consisting of the market-wage and hours-supply equations—the method of three-stage least squares applied to the model, for example. The reservation wage equation is estimated in a separate regression later, using the participation probit index as an additional right-hand side variable and the predicted wage as the dependent variable.[5]

This procedure uses all the information available and uses asymptotically efficient estimation methods relative to the information used in each stage but sacrifices some efficiency by separating the states. At each stage, therefore, only part of the information is utilized.

Thus far the above description of the methods applies to one given cross section. Accounting for the panel time-series aspects may be done simply by using levels as well as between-year differences of variables. A brief explanation of the method for two panels indicates the approach to be taken (which is equivalent in efficiency at the limit to full maximum likelihood estimates in a linear model). Let the hours equation in years 1 and 2 be:

$$H_1 = X_1\alpha_1 + Z\beta_1 + \epsilon + U_1$$
$$H_2 = X_2\alpha_2 + Z\beta_2 + \epsilon + U_2$$

where X are variables that change over time (incomes, wages, experience, children) and Z are fixed variables (race, past education, father's occupation and education, children ever born to older women, etc.) U_1 and U_2 are transitory residuals, and ϵ is a permanent individual residual component. The coefficients α and β are allowed, in the general case, to change between years but are the same for all individuals in each cross section.

Writing $X_2 = X_1 + \Delta X$ for the matrix X, allowing for a form of nonlinearity by substituting $(\alpha_{21}X_1 + \alpha_{22}\Delta X)$ for α_2X_2, and differencing the two equations gives:

$$\Delta H = (\alpha_{21} - \alpha_1)X_1 + \alpha_{22}\,\Delta X + (\beta_2 - \beta_1)Z + \Delta U.$$

Thus, testing for $\alpha_{21} - \alpha_1 = 0$ and $\beta_2 - \beta_1 = 0$ may determine the stability of the equations over time (for a given range of variables (X_1, Z)), while testing $\alpha_{22} = \alpha_{21}$, detects possible changes in parameters due to nonlinearities associated with changing the range of X from X_1 to X_2. The latter test is performed in a joint estimation by 3SLS of the first-year level equation and the difference equation. This procedure provides asymptotically efficient estimates, while allowing for a (fixed) covariance between the corresponding residuals $(\epsilon + U_1)$ and $(U_2 - U_1)$, (equal to $-\sigma_{U1}^2$).

If any of the tested coefficients are virtually equal, new estimates can be obtained using these equalities as additional restrictions. The covariance structure is also estimated consistently, thus identifying the permanent ver-

[4] See James J. Heckman, **Sample Selection Bias as a Specification Error,** Rand Corporation, 1976 (also published in **Econometrica,** January 1979); see also Giora Hanoch, **A Multivariate Model of Labor Supply: Methodology for Estimation**, Rand Corporation, 1976.

[5] Giora Hanoch, **ibid.**, pages 35-37.

sus the transitory components. Extension to 3 years is straightforward, using, for example, two difference equations ($H_2 - H_1$, $H_3 - H_2$) and one (H_1) level equation. This method can be extended to a simultaneous model with three or four equations for each year in an analogous way.

The identification of the actual segment on which each individual is located is determined by the available data—that is, whether he is retired, working and receiving full benefits, earning about M and receiving partial benefits (where his effective wage is $W(1 - t)$ and the effective nonwage income is $(Y + B + tM)$), or earning still more and receiving no benefits (with $W = W$ and $Y = Y$). As a control group, persons who are not eligible for the benefits (those, for example, who were not covered by the social security program in early jobs) are also included, with their expected supply curve H^* (W, Y, Z) throughout. The two-tiered tax system with $t_1 = 0.5$ and $t_2 = 1$ is treated in an analogous fashion.

Effects of current earnings on future benefits (particularly important for persons aged 62-64 subject to the early retirement provisions) can also be incorporated. Both the permanent nonwage wealth or income and the relevant current net wage are taken into account.

The discontinuity of supply at the point of entry is treated as in Hanoch[6] by specifying the reservation wage equation W_1^* $(Y;Z)$ separately from the basic supply function H^* $(W, Y; Z)$; the supply function applies conditionally for $W > W_1^*$ only (or $H^* > H_0$). Although W_1^* is not observed, it is estimated by using information on participation and on Z for the total sample and on wages of workers.

The labor suply in terms of annual weeks worked (a major form of variation when weekly hours are constrained by employers) can be specified as for H^*, but an additional modification is required [7] for the large proportion of individuals who work full year (52 weeks) and who are subject to a corner solution effect at that point. Given the estimated model, if any important and economically meaningful changes in parameters over time are detected, they may be incorporated into the model itself—by introducing additional interactions with age or time-trends in the coefficients.

Once a satisfactory version of the model is estimated, it is possible to derive from it the underlying preferences-structure in this population. It is also possible to use it to predict the labor-market behavior of any given individual (the "representative" individual, for example, whose residuals equal zero and whose other exogenous variables are equal to the population means) under various hypothetical conditions such as the reformed social security benefits structure or additional pension coverage.

With the estimates of the stochastic components taken into account, predictions with respect to aggregates (in the sample populations) may also be derived analytically for alternative simulated conditions or suggested policy reforms. Although in the existing data only two implicit tax rates on social security benefits applied (50 percent and 100 percent), the estimated equations and probabilities can be used, for example, to predict the aggregate effects on labor supply of changes in these tax rates or of their elimination (as in the current system in which only a 50 percent tax applies). In addition, predictions about the labor-supply behavior of future cohorts of the older population (not covered in the RHS data) may also be obtained if extrapolation of the observed trends in the exogenous variables and in the behavioral coefficients can be expected to remain valid.

Age, Cohort, and Period Effects

The analysis of data from studies such as the Retirement History Study, which combines longitudinal and cross-sectional information, has occupied an increasingly central role in a variety of social and economic studies. The discussion that follows deals with a focal problem in many of these studies: Estimation of the separate effects of time period, cohort of birth or vintage, and age. A strategy for dealing with the identification problem that arises because current year, year of birth, and age are linearly dependent is described briefly below.[8] Empirical results are presented that indicate the separate effects of age, cohort, and year on selected labor-market variables in the sample of older married men.

The usual cross-sectional pattern of the age profiles of labor-force participation or earnings compares persons of different ages in a given year. These profiles are influenced by differences between cohorts that result from factors other than aging. On the other hand, if fixed cohorts are observed over a period of time, the time-series profiles are confounded with cyclical effects and secular trends—also not related to age—in the economy and in the labor force.

It is now widely accepted that independent causal effects may be associated with each of the three factors. A current year may have cyclical effects, themselves the result of several underlying influences. The cohort, or year of birth, may have effects of its own: Individuals born in different years experience the same events at different ages, membership in an unusually large or small birth cohort may have important implications, etc. Economists in particular have been interested in the pro-

[6]See **ibid.**, and Giora Hanoch, **The Discontinuous Nature of the Supply of Labor** (paper presented at the French-Israeli meeting on human capital), Dijon, France, March 1979.

[7]See Giora Hanoch, **A General Solution for Estimating Period, Cohort, and Age Effects,** Center for Social Sciences, Columbia University, 1979.

[8]For a complete development, see **ibid.**, and Giora Hanoch and Marjorie Honig, **Age, Cohort, and Period Effects in the Labor-Market Behavior of Older Persons,** Center for Social Sciences, Columbia University, 1980.

files by age, since several aspects of economic behavior—labor-force participation, saving, consumption—are expected to exhibit life-cycle patterns.

Because of the effects of these three factors, which are different in the causal sense, any partial analysis involving only two of the factors may be biased. On the other hand, it is infeasible to specify a general unrestricted three-factor model in this case since the linear confounding that arises because age is the difference between year of birth and year of observation gives rise to underidentification, even after adding the three sets of effects. Although three separate causal effects may exist, not more than two effects may have independent linear components.

If a sufficient number of identifying restrictions are imposed on the parameters, however, the model is estimable. When the general additive model is just-identified (so that only the minimum number of necessary restrictions is imposed), the choice of the particular identifying restrictions does not affect the predicted values of the dependent variable; consequently, any measure of goodness of fit of the model to the observations is also invariant. The estimated separate effects of period, cohort, and age and their interpretation are, however, highly sensitive to the choice of restrictions.

To avoid this problem, the procedure adopted by most empirical studies in this field[9] has been:

(1) To choose the restrictions in the form of equality of effect of adjacent periods, cohorts, or ages; and

(2) To impose more restrictions than necessary and choose among the resulting overidentified models by inspecting the stability of the estimates and their goodness of fit.

These procedures are deficient on several grounds. First, the choice among alternative specifications requires a large and indefinite amount of computations and search, since the number of possibilities for equating adjacent effects is large and may be prohibitive with increased numbers of periods, cohorts, or ages and since the number of overidentifying restrictions imposed is determined arbitrarily and subjectively.

Second, no unique and generally accepted criteria for choice among estimates of given alternative specifications exist. Although goodness-of-fit criteria provide no clue whatsoever for choice among just-identified models, the redundant overidentifying restrictions may be tested statistically as nested hypotheses within the general

[9]This procedure is outlined in Kenneth O. Mason, William M. Mason, H. H. Winsborough, and W. Kenneth Poole, "Some Methodological Issues in Cohort Analysis of Archival Data," **American Sociological Review,** April 1973.

Table 1.—Employment of white married men measured by percent with social-security-covered earnings, by year and birth year

Year	Year of birth					
	1906	1907	1908	1909	1910	1911
Number in sample	516	624	584	611	628	720
1951	0.7848	0.7708	0.7945	0.8052	0.8136	0.7972
1952	.7868	.7788	.7945	.7970	.8152	.8027
1953	.8003	.7916	.8065	.8036	.8152	.8125
1954	.8003	.7852	.7996	.8117	.8121	.7958
1955	.8546	.8333	.8476	.8444	.8550	.8513
1956	.8682	.8509	.8664	.8674	.8789	.8611
1957	.8643	.8637	.8818	.8821	.8773	.8791
1958	.8720	.8573	.8732	.8837	.8837	.8763
1959	.8720	.8461	.8681	.8837	.8773	.8916
1960	.8701	.8461	.8869	.8919	.8710	.8805
1961	.8643	.8413	.8732	.9001	.8662	.8722
1962	.8624	.8349	.8715	.8903	.8630	.8819
1963	.8488	.8285	.8647	.8936	.8566	.8805
1964	.8507	.8237	.8630	.8837	.8455	.8680
1965	.8507	.8317	.8493	.8854	.8550	.8736
1966	.8468	.8301	.8544	.8805	.8566	.8722
1967	.8139	.8285	.8407	.8723	.8439	.8722
1968	.7655	.7996	.8253	.8739	.8407	.8652
1969	.7286	.7500	.8065	.8625	.8296	.8625
1970	.6976	.7019	.7568	.8477	.8216	.8500
1971	.5213	.6282	.6797	.7855	.8025	.8277
1972	.4089	.4791	.6095	.6988	.7308	.8013
1973	.3527	.3766	.4520	.6104	.6321	.7152
1974	.3081	.3317	.3441	.4631	.5509	.5972

Table 2.—Mean annual earnings of white married men in social-security-covered employment, by year and birth year[1]

Year	Year of birth					
	1906	1907	1908	1909	1910	1911
Number in sample	516	624	584	611	628	720
	Mean annual earnings					
1951	$2,983	$3,109	$3,080	$3,028	$3,016	$3,033
1952	3,124	3,159	3,152	3,177	3,131	3,037
1953	3,094	3,160	3,146	3,238	3,183	3,088
1954	3,040	3,186	3,110	3,096	3,131	3,075
1955	3,369	3,558	3,535	3,525	3,515	3,442
1956	3,550	3,620	3,653	3,581	3,578	3,572
1957	3,633	3,620	3,649	3,625	3,675	3,614
1958	3,560	3,585	3,617	3,630	3,603	3,562
1959	3,972	4,086	4,092	4,053	4,064	3,994
1960	3,944	4,072	4,096	4,047	4,067	4,082
1961	3,902	4,141	4,087	4,015	4,020	4,119
1962	3,952	4,200	4,155	4,115	4,083	4,142
1963	4,026	4,146	4,183	4,089	4,129	4,169
1964	4,086	4,153	4,218	4,210	4,222	4,258
1965	4,098	4,234	4,233	4,243	4,196	4,310
1966	5,144	5,486	5,363	5,447	5,406	5,525
1967	5,217	5,402	5,516	5,569	5,503	5,638
1968	5,776	6,021	6,150	6,217	6,252	6,412
1969	5,776	6,016	6,150	6,364	6,451	6,556
1970	5,178	5,792	5,788	6,283	6,361	6,529
1971	4,060	5,398	5,830	5,972	6,219	6,639
1972	3,924	4,399	5,642	6,353	6,386	7,015
1973	3,821	4,102	4,363	6,340	6,972	7,270
1974	3,923	4,122	4,519	4,929	6,675	7,925

[1] Excludes those with no covered earnings during the year.

model, and their earlier imposition may not be justified. Which restrictions in a given set should be viewed as redundant? No objective clear-cut indication is provided. Consequently, the resulting estimates remain to a large extent subjective, arbitrary, and open to dispute.

The procedure adopted in the present study proposes an alternative approach that is general, inexpensive, and straightforward. Briefly, since only two of the three linear effects of period, cohort, and age may vary independently, any two of these effects are specified in the equation explicitly in their linear forms. The remaining specific effects are independent of the chosen pair of linear effects and are treated symmetrically: For each of the underlying factors—time period, cohort, and age—the set of specific effects is purged of any linear trend by restricting them to be uncorrelated in the sample with their own corresponding time dimension, and all effects are conveniently measured as deviations from a particular year-cohort cell.

The interpretation of the specific effects under such restrictions is straightforward: The effects are analogous to regression residuals in each of the separate dimensions of time, cohort, or age, respectively, since the restrictions are analogous to the set of normal equations in a multiple linear regression. The identification problem is thus confined to the linear part of the model. If only linear age and linear period effects are specified in the equation, as in most of the following examples, the omitted linear cohort effect is embedded in the linear age effect that represents linear cross-sectional variation (in which age differences are indistinguishable from cohort differences). The linear period effect measures the trend of increase over time in excess of the cross-sectional age trend. Cohort effects need not be discarded entirely, however, since their nonlinear components, embedded in the deviations from the linear trend, are identifiable (and are independent of the arbitrary choice of the two linear effects).

An alternative solution to this identification problem

Table 3.—Employment of white married men measured by percent with social-security-covered earnings: Regression coefficients and adjusted and actual mean value of earnings, by year, birth year, and age effects[1]

Variable	Coefficient	Standard error of coefficient	Mean	
			Adjusted	Actual
Intercept	0.7972	0.0109	....	....
Linear:				
Age	-.0132	.0006	....	....
Year	.0034	.0006	....	....
Year:				
1951	0	....	0.7066	0.7944
1952	0.0011	0.0067	.7111	.7958
1953	.0124	.0070	.7257	.8050
1954	.0083	.0074	.7250	.8008
1955	.0539	.0077	.7740	.8477
1956	.0710	.0080	.7945	.8655
1957	.0799	.0084	.8068	.8747
1958	.0786	.0086	.8089	.8744
1959	.0772	.0087	.8109	.8731
1960	.0797	.0088	.8167	.8744
1961	.0765	.0088	.8169	.8696
1962	.0755	.0087	.8192	.8673
1963	.0727	.0086	.8198	.8621
1964	.0694	.0084	.8199	.8558
1965	.0745	.0082	.8284	.8576
1966	.0779	.0079	.8351	.8568
1967	.0725	.0076	.8332	.8453
1968	.0697	.0073	.8371	.8284
1969	.0705	.0070	.8379	.8066
1970	.0731	.0068	.8439	.7793
1971	.0514	.0066	.8256	.7075
1972	.0279	.0066	.8054	.6214
1973	-.0079	.0067	.7330	.5232
1974	-.0395	.0071	.7448	.4325
Year of birth:				
1911	0	....	.7954	.8370
1910	-.0051	.0031	.7903	.8206
1909	.0173	.0030	.8127	.8258
1908	-.0002	.0028	.7952	.7879
1907	-.0129	.0024	.7825	.7546
1906	.0082	.0017	.8036	.7539

See footnote at end of table.

Table 3.—Employment of white married men measured by percent with social-security-covered earnings: Regression coefficients and adjusted and actual mean value of earnings, by year, birth year, and age effects[1]—**Continued**

Variable	Coefficient	Standard error of coefficient	Mean	
			Adjusted	Actual
Age:				
40	0	......	.8883	.7972
41	.0245	.0139	.8996	.8082
42	.0283	.0134	.8902	.8110
43	.0296	.0132	.8783	.8006
44	.0405	.0131	.8759	.8065
45	.0512	.0130	.8734	.8163
46	.0615	.0135	.8704	.8301
47	.0707	.0137	.8664	.8424
48	.0870	.0139	.8694	.8596
49	.0969	.0140	.8661	.8715
50	.1035	.0141	.8590	.8720
51	.1122	.0142	.8549	.8716
52	.1264	.0142	.8559	.8748
53	.1308	.0141	.8470	.8677
54	.1393	.0140	.8423	.8659
55	.1460	.0139	.8358	.8625
56	.1554	.0138	.8320	.8614
57	.1568	.0136	.8201	.8519
58	.1672	.0134	.8172	.8521
59	.1702	.0132	.8070	.8458
60	.1735	.0130	.7971	.8354
61	.1682	.0129	.7785	.8119
62	.1302	.0128	.7273	.7506
63	.0807	.0127	.6645	.6731
64	.0465	.0129	.6171	.6193
65	-.0710	.0132	.4864	.4789
66	-.1426	.0137	.4012	.3765
67	-.1475	.0146	.3835	.3422
68	-.1648	.0170	.3529	.3081
Standard deviation of error	0.01090			
R^2	.9954			
Error degrees of freedom	88			

[1] Effects measured from year 1951, birth year 1911, and age 40 (OLS, unweighted). Assumes no linear time trend in year of birth effects.

may be provided if outside information is available regarding any one of the linear trends. In the following empirical analysis of annual earnings, for example, an estimate of the linear trend in earnings over time was based on an outside source and taken as given, and internal estimates of both the linear age and the linear cohort effects were thus provided.

Table 1 presents the employment rates of white married men (measured by the proportion with earnings covered by the social security program) by year and cohort.[10] Table 2 presents mean annual earnings in social-security-covered employment for those with some covered earnings during the year. Fixed ages appear on the diagonals from upper left. The usual cross-sectional view of the age profile appears in the rows, which compare employment in a given year for persons of different ages.

[10]Data are from social security records on earnings for a sample of white men aged 58-63 in 1969 (the first RHS interview) who were wage-earners and had a spouse present in 1969 and who remained in the study through 1973.

Table 4.—Mean annual earnings of white married men in social-security-covered employment: Regression coefficients and adjusted and actual mean value of earnings, by year, birth year, and age effects[1]

Variable	Coefficient	Standard error of coefficient	Mean	
			Adjusted	Actual[2]
Intercept	3033.0	108.0	......	......
Linear:				
Year[3]	185.87	......	......	......
Year of birth	-38.67	5.85	......	......
Age	-50.08	1.31	......	......
Year:				
1951	0	......	$3,062.3	$3,041.5
1952	-147.6	66.3	3,100.6	3,130.0
1953	-365.6	69.6	3,068.4	3,151.5
1954	-640.7	72.9	2,979.2	3,106.3
1955	-475.8	76.3	3,330.0	3,490.7
1956	-594.1	79.7	3,397.5	3,592.3
1957	-774.2	83.1	3,403.3	3,636.0
1958	-1037.4	85.1	3,326.0	3,592.8
1959	-803.2	86.5	3,746.1	4,043.5
1960	-1012.1	87.2	3,723.0	4,051.3
1961	-1229.9	87.2	3,691.1	4,047.3
1962	-1375.6	86.6	3,731.3	4,107.8
1963	-1563.4	85.4	3,729.3	4,123.7
1964	-1695.3	83.5	3,783.3	4,191.2
1965	-1863.7	81.2	3,800.8	4,219.0
1966	-863.6	78.5	4,986.7	5,395.2
1967	-937.5	75.5	5,098.7	5,474.2
1968	-383.0	72.4	5,839.1	6,138.0
1969	-405.7	69.4	6,002.3	6,218.8
1970	-627.5	67.1	5,966.3	5,988.5
1971	-660.0	65.8	6,119.7	5,686.3
1972	-395.2	65.6	6,570.4	5,620.0
1973	-157.1	66.8	6,994.3	5,478.0
1974	86.0	70.4	7,423.3	5,348.8
Year of birth:				
1911	0	......	4,586.2	4,792.0
1910	-69.2	31.1	4,478.3	4,659.9
1909	-52.5	29.9	4,456.4	4,547.8
1908	-60.0	27.4	4,410.2	4,388.6
1907	6.3	23.3	4,437.8	4,282.0
1906	-43.9	16.9	4,349.0	4,048.0

See footnotes at end of table.

These differences by age are affected by any cohort effects that may be present. Fixed cohorts may be observed in the columns in which cyclical and secular trends confound the pure age effect.

Table 3 contains regression coefficients (ordinary least squares unweighted) and adjusted means by age, cohort, and year based on the data in table 1 and the methodology described above. Effects are measured from 1951, birth year 1911, and age 40. The adjusted means are derived from the regression coefficients, corrected for the difference in mean level between the adjusted and actual series; actual means are unweighted means across cells in table 1.[11]

Under the assumption of no linear time trend in cohort effects, the linear age coefficient may be interpreted as a

[11]The actual mean for a given age, for example, is the unweighted mean along the diagonal.

Table 4.—Mean annual earnings of white married men in social-security-covered employment: Regression coefficients and adjusted and actual mean value of earnings, by year, birth year, and age effects[1]—**Continued**

Variable	Coefficient	Standard error of coefficient	Mean	
			Adjusted	Actual[2]
Age:				
40	0	......	$4,290.6	$3,033.0
41	78.4	137.5	4,318.9	3,026.5
42	213.9	132.6	4,404.3	3,082.3
43	359.1	130.6	4,499.5	3,128.8
44	448.3	129.7	4,538.6	3,214.4
45	501.7	129.3	4,541.9	3,245.2
46	601.6	134.1	4,591.7	3,351.8
47	701.1	136.0	4,641.2	3,438.8
48	778.3	137.6	4,668.3	3,578.8
49	861.2	139.0	4,701.1	3,735.7
50	953.1	139.9	4,742.9	3,837.7
51	1029.7	140.4	4,769.5	3,919.8
52	1107.1	140.4	4,796.8	4,001.5
53	1185.0	140.0	4,824.6	4,105.5
54	1263.9	139.2	4,853.4	4,143.5
55	1329.2	138.0	4,868.7	4,369.3
56	1375.9	136.5	4,865.3	4,600.5
57	1436.4	134.8	4,875.7	4,962.3
58	1488.7	133.0	4,877.9	5,343.3
59	1547.8	131.1	4,887.0	5,716.2
60	1505.2	129.3	4,794.3	6,010.0
61	1432.4	127.8	4,671.4	6,151.0
62	1216.8	126.6	4,405.7	6,201.3
63	1242.9	125.9	4,381.7	6,441.3
64	622.8	128.3	3,711.6	5,846.6
65	-885.1	131.0	2,153.6	4,437.8
66	-1295.9	135.7	1,692.7	4,181.7
67	-1663.2	144.7	1,275.3	3,971.5
68	-1831.7	168.6	1,056.8	3,923.0
Standard deviation of error	108.04			
R^2	0.9947			
Error degrees of freedom	88			

[1] Effects measured from year 1951, birth year 1911, and age 40 (OLS, unweighted). Excludes those with no covered earnings during the year.

[2] Unweighted means across cells.

[3] Linear trend over time was taken as given, based on linear time trend (relative to the mean), from an "Index of Weekly Wages in Manufacturing," **Employment and Earnings**, Bureau of Labor Statistics, table C-1, 1977.

pure age trend. A decline of .0132 per year of age is shown in the proportion with covered earnings. The specific age coefficients, which measure deviations from the linear trend, indicate a nonlinear pattern of rising then rapidly falling employment rates. The linear time trend shows a small increase (.0034) per year that may in part reflect the increase in social security coverage during this period. The specific year effects indicate cohort differences in employment around a constant mean.

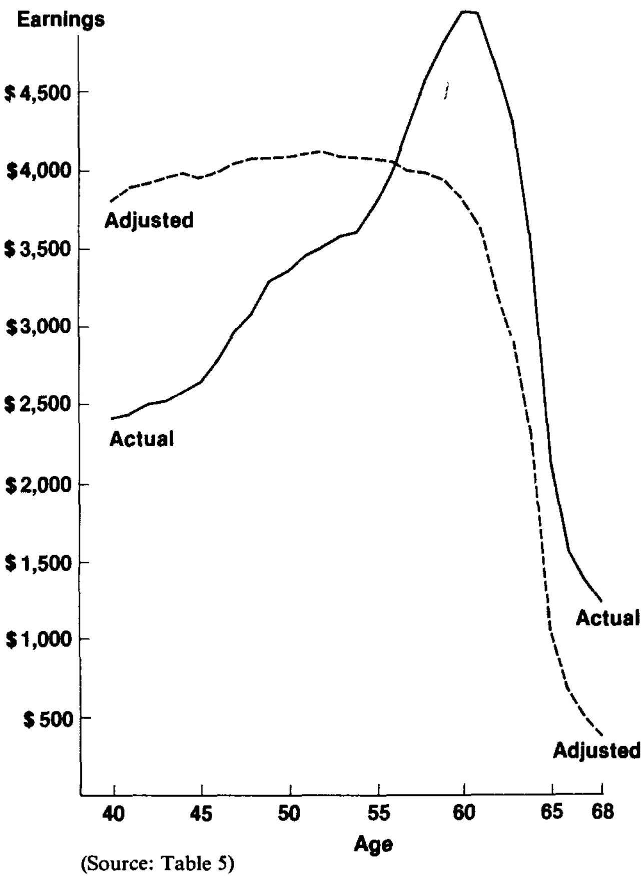
Table 4 presents estimates of linear and specific age, cohort, and year effects in annual earnings in covered employment (including those without covered earnings during the year) based on the data in table 2. The linear trend over a period of time is taken as given, based on the linear trend in an economy-wide index of weekly wages in manufacturing. This outside information permits identification of linear cohort and linear age effects. The linear cohort trend is negative, implying a decline in annual earnings for older cohorts—ranked from youngest (1911 = 0) to oldest—that may reflect lower levels of education for successively older cohorts. The linear age effect is also negative and represents an average decline of $50 in annual earnings per year of age. The specific age effects are similar to those found for employment in table 3, rising with age then declining sharply in the latest years.

Table 5.—Actual and adjusted mean value of annual earnings of white married men in social-security-covered employment, by age, year, and birth year[1]

Variable	Mean	
	Adjusted	Actual[2]
Age:		
40	$3,811	$2,418
41	3,885	2,446
42	3,921	2,500
43	3,952	2,505
44	3,975	2,595
45	3,967	2,656
46	3,997	2,791
47	4,021	2,904
48	4,059	3,084
49	4,072	3,257
50	4,074	3,347
51	4,077	3,418
52	4,106	3,500
53	4,086	3,562
54	4,088	3,587
55	4,069	3,771
56	4,048	3,965
57	3,999	4,237
58	3,986	4,558
59	3,944	4,835
60	3,822	5,020
61	3,637	4,993
62	3,204	4,645
63	2,912	4,302
64	2,290	3,598
65	1,048	2,120
66	679	1,568
67	489	1,358
68	373	1,209
Year:		
1951	2,164	2,416
1952	2,205	2,491
1953	2,227	2,537
1954	2,386	2,487
1955	2,577	2,959
1956	2,699	3,109
1957	2,746	3,181
1958	2,690	3,142
1959	3,038	3,530
1960	3,041	3,543
1961	3,015	3,519
1962	3,057	3,563
1963	3,057	3,555
1964	3,102	3,587
1965	3,149	3,619
1966	4,164	5,395
1967	4,248	4,630
1968	4,888	5,091
1969	5,029	5,029
1970	5,035	4,693
1971	5,052	4,108
1972	5,292	3,647
1973	5,407	3,057
1974	5,529	2,471
Year of birth:		
1911	3,648	3,966
1910	3,539	3,765
1909	3,622	3,723
1908	3,507	3,429
1907	3,473	3,217
1906	3,495	3,045

[1] Includes those with no covered earnings during the year.
[2] Unweighted means across cells.
Source: Tables 3 and 4.

Chart 3.—Actual and adjusted mean annual earnings of white married men in social-security-covered employment, by age

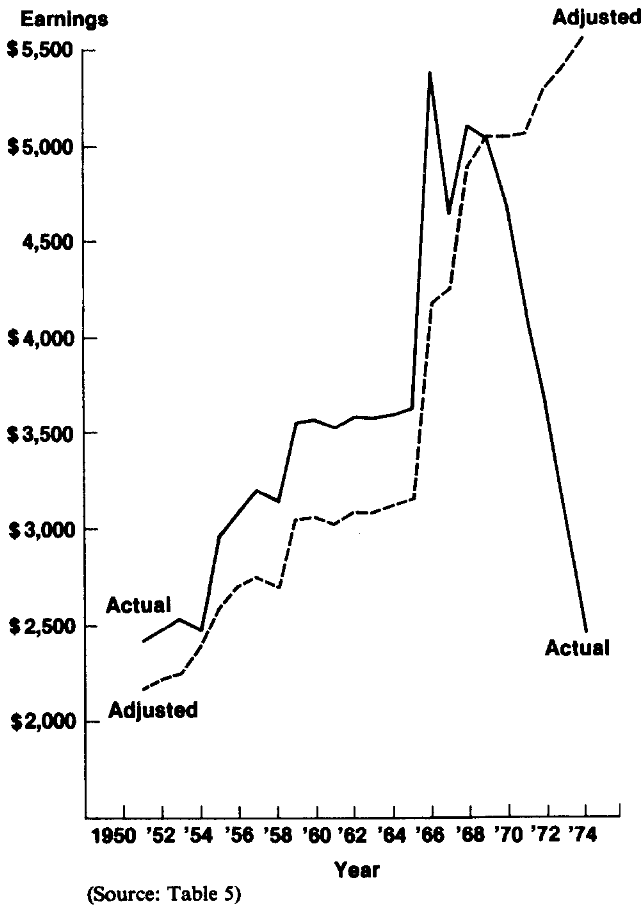


(Source: Table 5)

Cohort effects exhibit a somewhat different nonlinear pattern from those relating to employment, and cyclical movements are more pronounced than those appearing in specific year coefficients in table 3.

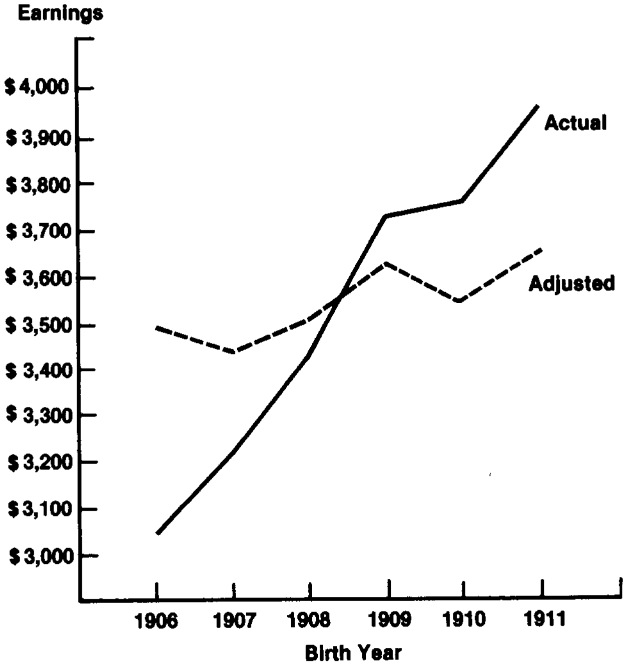
Table 5 presents actual and adjusted mean values for annual earnings in social-security-covered employment (includes those without covered earnings during the year). These data reflect changes in both labor-force participation and mean annual earnings of workers and are derived from the estimates in tables 3 and 4. Charts 3-5 present these results graphically. Although the actual mean earnings by age shown in chart 3 also reflect cohort and year differences, the adjusted means provide an age profile of earnings unclouded by either of these effects. A more reasonable and smoother age profile is obtained with earnings rising slightly through the middle years, leveling off in the fifties, and beginning to decline in the late fifties. A substantial decline is observed at age 62—the age of eligibility for early retirement benefits under the social security program—and a larger decrease at age 65, the standard retirement age.

Chart 4.—Actual and adjusted mean annual earnings of white married men in social-security-covered employment, by **year**



(Source: Table 5)

Chart 5.—Actual and adjusted mean annual earnings of white married men in social-security-covered employment, by **year of birth**



(Source: Table 5)

Adjusted earnings by year (chart 4) show a steady rise with productivity and price increases, compared with the series of actual means, by year, that rise then fall as a result of aging in the sample. The adjusted means by cohort (chart 5) are fluctuating around the overall mean in annual earnings, compared with the trend in the actual means, which again reflects a bias because of an aging trend.

Table 6 presents age, year, and cohort effects in selected labor-market variables from the RHS panel data: Employment at time of interview, weekly hours of work on current job, hourly wage, and health. Data from the three interview years 1969, 1971, and 1973 are used, and ages of the sample range from 58 to 67. All effects are measured from 1969, birth year 1911, and age 58. As in table 3, no linear trend in cohort effects is assumed. Modified equations are presented when the omission of insignificant variables does not result in a significant reduction in explained variance (F-test statistics for comparison of basic and modified equations appear at the bottom of the table).

The modified employment equation indicates a negative linear age trend and a fluctuating pattern of specific age effects around the trend. The linear time trend is also negative, reflecting the general downturn in the economy in this period. No specific cohort differentials are evident.

The health measure is a dummy variable indicating whether the respondent reported that poor health restricted the amount or type of work he was able to do. The modified equation indicates a positive age trend, with the proportion reporting a health limitation increasing by 2 percent per year of age. The combined linear and specific year effects show a decline from 1969 to 1971 and a larger increase to 1973 in the proportion reporting health problems.

The negative linear age effect in weekly hours of work on the current job (modified equation) indicates a decline of slightly more than 1 hour per year of age. The specific age coefficients show a rising then falling pattern around this trend. Neither the year effects nor the specific cohort effects are significant.

The basic equation of the log of hourly wage indicates nearly all effects are significant. The pure age trend shows a 2.5-percent decline in wages per year of age and, like weekly hours of work, a rising then falling pattern in the nonlinear component. Significant cohort differences are observed, with the 1910 and 1906 cohorts receiving relatively lower hourly wages.

The age, cohort, and period effects indicated in the preceding tables provide a building block in the model of the labor-market behavior of older persons. The methodology described above allows for separation of each of these effects and determination of their statistical significance. Those found to be unimportant in the determination of any of the labor-market components may be ignored. Significant effects are likely to represent a number of underlying factors that may then be examined. Age effects, for example, would be expected to be correlated with human capital and other variables in the more general context of labor-supply behavior. A detailed quantitative analysis of these variables—and the study of additional issues involving age, cohort, and period effects—is currently in progress.

Table 6.—Modified and basic equations for selected labor-market variables of white married men with social-security-covered earnings, by age, birth year, and year effects,[1] panel years 1969, 1971, 1973[2]

Variable	Employment [3]		Health limitation		Weekly hours of work [4]		
	Basic	Modified	Basic	Modified	Basic	Modified	Basic
Intercept	[5] 2.5530	[5] 2.5530	[5] 0.2417	[5] 0.2304	[5] 43.2374	[5] 43.2374	[5] 1.2300
Linear:							
Age	[5] -.3808	[5] -.3784	[5] .0208	[5] .0203	[5] -.8987	[5] -1.0330	[5] -.0246
Year	[5] -.1146	[5] -.1197	[5] .0117	[5] .0130	[5] -.2774	-.1356	.0549
Age:							
59(1)	-0.0524	-0.0496	0.0045	....	0.6559	0.7724	[5] 0.0491
60(2)	[6] .1994	[5] .2241	-.0144	....	[6] 1.5588	[5] 1.6220	[5] .0309
61(3)	[6] .2494	[5] .2171	-.0160	....	[5] 2.9105	[5] 3.1542	[5] .0766
62(4)	.0776	.1111	-.0216	....	[5] 3.4613	[5] 3.4835	[5] .0905
63(5)	[5] .2622	[5] .2387	-.0139	....	[5] 3.9360	[5] 4.1174	[5] .0723
64(6)	[6] .2354	[5] .2856	.0048	....	[5] 3.7593	[5] 3.7927	[5] .0838
65(7)	[6] -.1554	[5] -.1920	.0038	....	[6] 1.7737	[5] 1.9867	[5] .0605
66(8)	.0680	.1191	-.0061	....	.2022	.1425	[5] .0257
76(9)	[6] .2485	[5] .2428	[6] -.0285	-0.0216	-.6100	-.5532	[5] -.0135
Year of birth:							
1910(1)	.0052	....	.0021	....	-.0179	....	[5] -.0479
1909(2)	0523	....	[6] -.0326	[5] -.0322	-.3643	....	[5] -.0148
1908(3)	-.0620	....	.0169	.0194	[6] -.9052	-.6511	[5] -.0145
1907(4)	.0503	....	-.0093	....	-.5672	....	.0004
1906(5)	-.0041	....	-.0030	....	.4378	1.0229	[5] -.0291
Year 1971(2)	[5] .1289	[5] .1325	[5] -.0267	[5] -.0278	.2369	....	[6] -.0042
Standard deviation of error	0.03912	0.0477	0.00734	0.0132	0.31878	0.3199	0.00153
R^2	[5] .9998	[5] .9993	.9980	[5] .9767	.9979	[5] .9965	[5] .9999
Error degrees of freedom	6		11		5		
F-test (vs. basic)	2.0		2.7		1.0		
$F_{.95}$ (d-3,3)	9.3		8.9		9.6		

[1] Effects measured from year 1969, birth year 1911, and age 58 (OLS, unweighted).

[2] Three degrees of freedom for error (three restrictions on parameters). Assumes no pure year of birth trend.

[3] Logit transformation of means of dummy variable (log P/1-P).

[4] Excludes those with no covered earnings during the year.

[5] Significant at 1-percent level of confidence.

[6] Significantly different from 0 at 5-percent level of confidence.