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We use the Current Population Survey to describe what we believe are the most salient aspects of labor force behavior of older men and women during the last 2 decades. First, we show that early retirement has increased dramatically, and this trend continued through the 1980s. Second, we show that the factors that most sharply distinguish propensities toward early retirement are those usually associated with low wages. Third, we show that trends in reduced participation for older men parallel those for younger men, while a pattern of increasing female participation is to be expected given the behavior of younger cohorts.

I. Introduction and Overview

We use the matched and unmatched March Current Population Survey (CPS) to describe what we believe are the most salient features regarding the labor force behavior of men and women aged 49 and over during the last 2 decades. Our primary interest is in summarizing data over a relatively

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long period to question the conventional wisdom regarding causes of declining retirement age and provide the stylized facts with which future research must contend.

The matched CPS is a little used but very rich source of information on dynamic labor force behavior and consists of a long sequence of short panels constructed by matching individual records across adjacent-year files of the March CPS. The use of these data is one of the novelties of this article. Unlike earlier studies, we are in the unique position of being able to analyze participation and labor force transitions over a long period of time and for many birth cohorts. We concentrate equally on men and women, although we abstract from the difficult problem of how decisions are made within a household.

Our data begin in 1968 when male participation rates appear to be at a post–World War II high. Since then, male participation (defined as weeks employed out of 52) has fallen at all ages, but most abruptly for men aged 62–65.[1] In 1968, participation was 91.5% at age 50, 83.8% at age 60, 58.4% at age 65, and 28.5% at age 69. During the next 20 years, the participation rate fell by about 4–5 percentage points for men aged 50, and the drop was progressively larger for older men up through age 64. For men aged 60, the 1968–89 drop in participation was 15.8 percentage points, and it was 30.5 percentage points (to a level of 36.6%) for men 64 years old. For men aged 67–69, participation rates by the end of our data were roughly 12 percentage points lower than at the beginning.[2]

For women, the 1968 and 1989 age-participation profiles intersect at age 60. Women younger than 60 have higher participation rates now than they did 20 years ago. Women aged 60 and over have lower participation rates, although the reduction is only 3–4 percentage points for women aged 62 and older.

Among those aged 50–62, the male-female difference in participation rates narrowed by roughly 15 percentage points during the 2 decades. The narrowing is more pronounced at the modal retirement ages 62–64, where participation rates for men dropped 20–25 points more than for women. In the main part of this article, we examine trends in transitions between labor force states—full-time, part-time, and retirement (nonparticipation)—rather than levels of participation per se. After a brief description of the data and the operational definitions of labor force status in Sections II and III, we provide nonparametric summaries of age-specific transition

[1] Weeks worked are approximate full-time weeks, with weeks for those who usually worked less than 35 hours reduced by one-half. Detailed tabulations used in this study of labor force participation and transitions by sex, age, and year are available from us on request.

[2] These calculations are from the unmatched March CPS. The 1968 data are simple averages of age-specific participation rates for 1967, 1968, and 1969 (from the 1968–70 surveys). The 1989 data are from the 1989–91 surveys.

rates through time in Section IV. This section shows that retirement at the Social Security early retirement age (62) has increased dramatically, and this trend continued through the 1980s. It also shows that exit rates from full-time work tend to peak during recessions for older men just as they do for younger men.

Labor force transitions are the natural metric for studying retirement behavior, and it is not surprising that most existing studies adopt this approach to the data. It is important, however, to emphasize that labor force participation rates *per se* are convolutions of transitions that have accrued over the full careers of the individuals we observe. At the end of Section IV we provide an accounting linking changes in cross-section age-participation profiles to changes in transition rates.

Section V uses a multinomial logit specification to parameterize age and time effects and to analyze the effects of other personal characteristics, such as race, education, and family composition. We find that the factors that most sharply distinguish propensities toward early retirement are those usually associated with low wages. Since wages and wealth are presumably correlated, we suspect that increasing wealth is not the most important contributor to falling ages of retirement. It is not that we believe that attributing declining ages at retirement to increased pension and Social Security wealth is misguided, but the fact is that those who retire earliest have characteristics that are usually associated with low wages and low wealth, and over the time span of our data, as wage dispersion has increased, so has dispersion in retirement rates. As an example, education has always been a good predictor of labor force exit rates for older men and women, but this is much more important today than 20 years ago.

Section VI changes the format, by switching from transitions to participation rates. We show in this section that trends in reduced participation for older men parallel those for younger men: with respect to the time profile of participation, one age is much like another. Women are different. Although there is evidence of leveling in participation rates of women aged 60 and older, the fact is that participation of these women has changed only trivially during the last 25 years. Almost all of the action in increasing participation has occurred at younger ages. Because participation has increased for younger women, those who will reach advanced ages in the future will have had very different work histories from those experienced by the older women that we observed to date.

Our purpose in this section is to suggest that the search for explanations of trends in the labor force behavior of older people should primarily emphasize the larger question surrounding participation in general, and only secondarily should the peculiarities of advancing age be addressed. We close on this point because we believe that the retirement literature is too specialized. Obviously, old age has its own distinguishing aspects, but it seems that the major trends in the data cannot be attributed to them.

Finally, Section VII contains the summary, some speculations, and a few suggestions for further research.

II. The Data

Our data come from two CPS sources: matched and unmatched March files. The matched files are used in Sections IV and V to study transitions between labor force states. The unmatched files are used in Section VI to compare trends in labor force participation of men and women of different ages.

Although the Census Bureau releases only unmatched files, it is possible to match individual records across adjacent surveys. The CPS uses a "4 in—8 out—4 in" monthly rotation scheme; that is, respondents are first interviewed for 4 successive months, and then, after an 8-month lapse, they are interviewed for an additional 4 months. For each rotation group the second wave of four interviews occurs 1 year later in the same calendar months as the first wave. Hence, if the sample size was constant and there was no attrition, half of sample in the March survey of any given year could in principle be matched with the following year's March survey, while the other half could be matched with the previous year's March survey. Because of attrition and the addition of new households, this ideal cannot be attained. As a general rule, in any given survey about three-quarters of the potential household matches can be realized. Individual match rates within matched households average to about 90%.[3]

When we restricted attention to civilians aged 49–68 in the first survey, the sample in a matched CPS file consisted on average of about 5,000 men and 5,700 women for each available year. This is comparable in size to the Retirement History Survey (RHS) and the National Longitudinal Survey (NLS) of Labor Market Experience of Older Men, the two data sets most commonly used to study retirement behavior.[4]

Matched CPS data offer several advantages over the RHS and the NLS. First, the CPS provides direct information on labor force transitions over a 1-year interval, as opposed to the 2-year interval of the RHS and NLS. Second, unlike the RHS and NLS that follow a single cohort, our sample spans several cohorts and maintains a distribution by age and other individual characteristics that is relatively stable, changing only with sample noise and with aggregate movements. Third, our data cover the period

[3] The March–March files from 1968 to 1976 were matched by Katz, Teuter, and Sidel (1984). Alan Pitts (1988) matched the 1977–79 files using modified Katz et al. scores. The 1979–91 files were matched by Finis Welch. A description of his algorithm can be found in Welch (1993).

[4] The RHS is a panel of 11,153 individuals (8,131 men and 3,022 unmarried women) aged 58–63 in 1969, interviewed every 2 years from 1969 to 1979. The NLS–Older Men is a panel of 5,020 men aged 45–59 in 1966, interviewed 12 times between 1966 and 1983.

from 1968 to 1990, although we have only 19 of the 23 potential matches. Household identifiers were scrambled in 1972 and shifted in 1977 and 1986. As a result, matched files for 1971–72, 1972–73, 1976–77, and 1985–86 are not available.

A disadvantage of the matched CPS is that we can only analyze changes over a 1-year period. More complicated patterns of state dependence cannot be addressed.

Another problem is high attrition. The main sources of attrition in the CPS are individual and household mobility, noninterview, and temporary addition of Hispanic households to the March survey.[5] Although lower than for the CPS as a whole, individual attrition rates for people aged 49–68 in the first survey average to about 20%.[6] Since labor force behavior and migration are correlated, transition estimates based on the matched CPS may be biased. In Peracchi and Welch (1993) we show that the matching process induces biases in estimated participation rates, but we find no evidence of systematic bias in the estimates of labor force transitions.

III. Labor Force Status

An operational definition of retirement is complicated by the fact that retirement has two distinct but related aspects. The first involves a decision about when and how much to reduce the amount of labor supplied to the market. The second involves deciding when to start withdrawing annuities from accumulated pension or Social Security wealth. The labor participation aspect leads to identifying retirement either with withdrawal from the labor force or with a sudden and sizable reduction in labor supply. The other aspect leads to identifying retirement with recipiency of pension or Social Security income.

In this article we focus on the labor force participation aspect of retirement. We use an instantaneous definition of labor force status, based on the number of hours worked during the week preceding the survey, integrated with some retrospective information, to define three mutually exclusive labor force states: full-time participation, part-time work or partial

[5] Pitts (1988) analyzes the distribution of unmatched households by sources of nonmatch. He finds that, on average over the period 1979–83, 42% are "movers," 29% are noninterviews, 12% are oversampled Hispanics, while for the remaining 17% the reason cannot be discerned using the public release tapes. The nonidentifiable sources include other additions or deletions of households by the Census for purposes of sample design and errors in transcribing households or individual identifiers onto tape.

[6] By comparison, individual attrition rates in the male sample of the RHS are 13.2% at the time of the second survey and 47.1% at the time of the last survey. Individual attrition rates in the NLS–Older Men are 5.6% at the time of the second survey and 47.4% at the time of the last survey.

retirement, and nonparticipation or full retirement.[7] Specifically, a person is a full-time worker if he or she is at work and either is working at least 35 hours or, if working less than 35 hours, usually works 35 hours per week or has a full-time workweek of less than 35 hours. One who has a job but did not work in the week preceding the survey is classified as full-time if the job usually requires more than 35 hours per week. Those unemployed are classified as full-time workers if they are looking for a full-time job, have been unemployed for not more than 1 year, and worked at least 2 weeks at some time during the previous 5 years.

Those who are at work or have a job are classified as part-time if they are not full-time. Similarly, an unemployed person is classified as part-time if he or she is not full-time. Finally, one who is neither full-time nor part-time is classified as retired (out of the labor force).[8]

Table 1 presents adjacent years transitions between labor force states for the pooled surveys. Among the men aged 49–68, 68.5% participate full-time when first observed, and about 64% are full-time 1 year later. Similarly, 27% are retired at the time of the first survey, and 31% are retired at the time of the second survey. The fraction of men who are part-time does not change much between the two surveys. Of the men who are full-time in the second survey, 96.6% were full-time in the first, 1.9% were part-time, and 1.5% were retired. For women, table 1 shows a 1-year decline in full-time participation from 33% to 31%, with an associated increase in the fraction retired from 57% to 59%.

Transitions to and from part-time participation are important. Roughly one-fourth of employed women are part-time, yet transitions from part-time account for 45% of all entries into retirement. The contrast is even sharper for men, where part-time accounts for one job in 15, and yet transitions from part-time represent nearly one-fourth of the entries into retirement.

IV. Labor Force Transitions: Nonparametric Analysis

We analyze trends in labor force behavior within the context of a simple three-state Markov chain model linking labor force state probabilities in 2 successive years. There are several reasons for doing so.

[7] A spot definition results in higher frequency of transitions than one based on averages of periods or retrospective data, but it appears preferable to one based on retrospective information because it reduces the probability of recall errors, rounding, heaping, etc.

[8] We found no major inconsistencies between our instantaneous definition of labor force status and the results from retrospective questions. For example, about 86% of the men and 75% of the women classified as full-time in both surveys worked at least 50 weeks in the year before the second survey, usually at least 36 hours per week. Of the men classified as part-time in both surveys, about 60% worked for at least 40 weeks, usually between 6 and 30 hours per week, whereas only 15% worked less than 27 weeks. For women, these percentages are 70% and 9%, respectively. Finally, 82% of the men and 96% of the women classified as retired in both surveys worked 0 weeks the year before the second survey.

Table 1
Adjacent Year Distribution by Labor Force Status: All Surveys 1968–91, All Ages 49–68

First Year	Second Year FT	PT	RT	Total	Sample Size
Men:					
FT	90.4	3.0	6.6	100.0	65,803
	61.9	2.1	4.5	68.5	
PT	25.5	46.8	27.7	100.0	4,746
	1.2	2.3	1.4	4.9	
RT	3.6	3.4	93.0	100.0	25,546
	1.0	.9	24.7	26.6	
Total	64.1	5.3	30.6	100.0	96,095
Women:					
FT	83.5	7.3	9.2	100.0	35,619
	27.4	2.4	3.0	32.8	
PT	19.2	57.9	22.9	100.0	11,353
	2.0	6.0	2.4	10.4	
RT	2.3	3.0	94.7	100.0	61,714
	1.3	1.7	53.8	56.8	
Total	30.7	10.1	59.2	100.0	108,686

NOTE.—FT = full-time work, PT = part-time work, RT = retired.

First, a Markov model provides one way of distinguishing between two types of information contained in the CPS. One is the "stock" information contained in the unmatched files, which is the basis for estimating participation probabilities. The other is the "flow" information contained in the matched files, which is the basis for estimating transition probabilities between labor force states.

Second, most structural models of retirement behavior, such as the ones of Rust (1989) and Berkovec and Stern (1991), are formulated in terms of transition probabilities implied by a stochastic dynamic discrete choice problem. The reduced-form estimates obtained from our data may be viewed as the stylized facts that a structural model ought to be able to reproduce.

Third, since participation probabilities are essentially convolutions of transition probabilities, the existence of a stable relationship between personal characteristics and transition probabilities guarantees the existence of a stable relationship between personal characteristics and participation probabilities. In this case, a Markov model represents a simple and useful tool for predicting future participation given current labor force status.

A. Age-Transition Profiles

If π is the vector of state probabilities, the link between state probabilities in 2 successive years for a group of persons of age a at time t is given by

$$\pi(a + 1, t + 1) = \Lambda(a, t)^T \quad \pi(a, t), \tag{1}$$

where $\Lambda(a, t)$ is the 1-year transition probability matrix, whose generic element $\lambda_{ij}(a, t)$ is the conditional probability of moving from state i to state j over a 1-year period.[9] A simple nonparametric estimate of $\lambda_{ij}(a, t)$ is its sample analog, namely, the fraction of individuals in state i at age a who move to state j between year t and year $t + 1$.

In figure 1 we present nonparametric estimates of the average matrix of transition probabilities by sex and age at the time of the first survey. The estimates are weighted averages of observed transition frequencies for cells defined by sex, age, and year, with weights equal to the marginal distribution by sex, year, and labor force status in the first survey. Since the age profiles of transition rates have features that are well known to retirement analysts, we make only a few remarks.

The first row of the matrix corresponds to the exit from full-time work. Although women have higher exit rates than men, especially at earlier ages, the profiles of the hazards are otherwise quite similar. The hazard of exit into retirement increases slowly at earlier ages, rises rapidly after age 59, has a first spike at age 61, a second sharper spike at age 64, and then declines somewhat. Jumps at age 61 and 64 can also be found in the hazard of exit into part-time work. The jumps or spikes at ages 61 and 64 have typically been explained by the presence of convex kinks in the lifetime budget constraint facing a worker. The spike at age 64 has been explained by current Social Security regulations, which discourage work beyond age 65 and whose effect may be reinforced by private pension provisions and, prior to 1983, by mandatory retirement rules. The spike at age 61 has been explained by the presence of liquidity constraints in the credit market, coupled with the Social Security early retirement option at age 62 (see, e.g., Kahn 1988). Incentive effects induced by private pension plans do not seem to be an explanation in this case.

Except for sampling noise, the hazards on the second and third row appear to be smooth, monotonic functions of age. Compared with full-time work or retirement, part-time work is a highly transient state, especially for men. The probability of remaining in a part-time job over a 1-year period shows very little age dependence for women but increases significantly with age for men.

Exit rates from retirement are low, especially for women, and decline steadily with age. After age 60, retirement is almost an absorbing state with exit rates below 5% for both men and women.

[9] It is understood that when first-order Markov transition matrices are formed from group averages, they need not describe individual participation profiles, due both to heterogeneity and state dependence (see, e.g., Heckman and Willis 1977), but they definitionally describe the average behavior of the group used to form them.

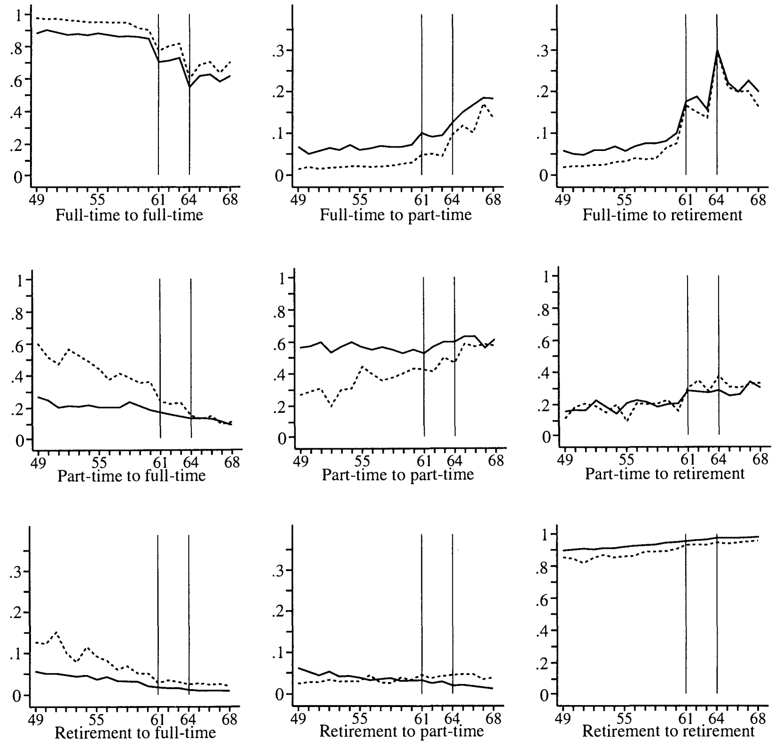

FIG. 1.—Transition rates by age

B. Time Trends

We now turn to the time-series evidence to verify the existence and nature of trends in the transitions between full-time work, part-time work, and retirement over the period 1968–90.

In figure 2 we present nonparametric estimates of the average matrix of transition probabilities by sex and year. Cell-specific transition rates are now averaged by age, with weights equal to the marginal distribution by sex, age, and labor force status in the first survey. To help in identifying trends in the data, we also plot a smooth scatter plot obtained by fitting a locally linear trend to the estimated transitions using Cleveland's (1979) loess method.

The figure shows a clear upward trend in the exit rates from full-time into part-time work, for both men and women. Exit rates from full-time work into retirement follow an upward trend up to the early 1980s, at least for men, and show some evidence of a reversal after 1982. For men, we also observe a negative trend in exit rates from part-time work and

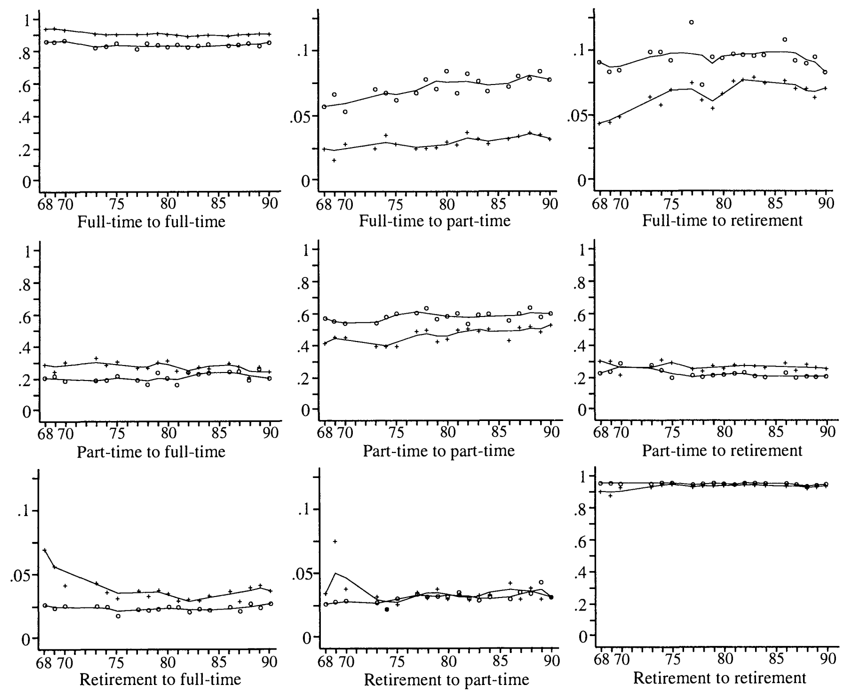

FIG. 2.—Transition rates by year

retirement into full-time work, and a positive trend in part-time retention rates. For all other transition rates, no clear trend emerges.

The analysis of trends is complicated by the presence of a cyclical component in exit rates from full-time and part-time work. The pattern for full-time work appears to be inversely related to the business cycle, with peaks during the 1974 and 1980–82 recessions. Two effects are likely to be at work here and act in the same direction. First, many firms are known to offer special early retirement during recessions to induce voluntary reduction of the workforce. Second, labor force participation may respond to falling real wages.

The next three figures show that trends in exit rates from full-time work have not been uniform at all ages. Figure 3 compares exit rates from full-time work for four age groups, namely 49–53, 54–60, 61–63, and 64–68. For both men and women aged 61–63, exit rates into retirement show a clear upward trend that continues throughout the 1980s. For people aged 64 and older, there is instead some evidence of a downward trend after the 1980–82 recession.

Figure 4 compares the hazards of exit from full-time work for the two subperiods 1968–77 and 1984–90. What captures the eyes is the behavior of the hazard for exit into retirement, which is characterized by a sharpening

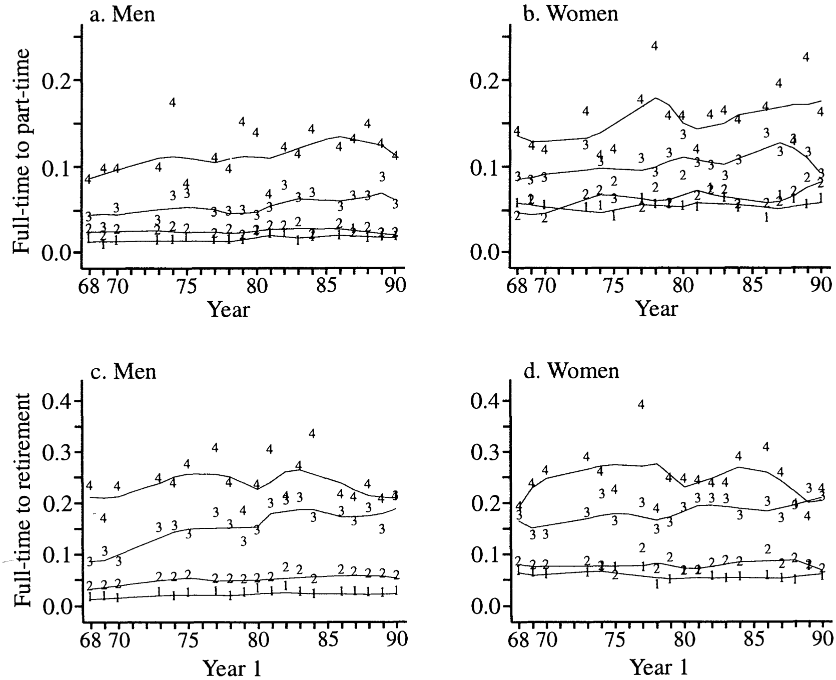

FIG. 3.—Transition rates by year and age group

of the age 61 spike, and a corresponding decline of the age 64 spike. The emergence of the spike at age 61 is more striking for men, while the decline of the spike at age 64 is more pronounced for women. For women, a downward shift in the hazard before age 57 is also noticeable.

We interpret the emergence of the spike at age 61 as an indication that the number of people who would have retired earlier, if they had the possibility to do so, has increased over time. It is hard to suspect a role for Social Security. First, the ability to retire at age 62 has existed since 1963. Second, changes in Social Security, starting with the 1977 reform, have gone in the direction of discouraging early retirement. Perhaps one ought to reassess the incentive effects of those changes. If credit market imperfections are the explanation for retirement at age 62 then, in order to understand why this has been rising steadily through time, one should concentrate on factors that explain why an increasing fraction of workers are liquidity constrained.

Figure 5 provides further detail by showing how the hazards for exit from full-time work have changed by birth cohort. We consider five equally spaced birth cohorts, spanning the years from 1907 to 1931. The first cohort, those born between 1907 and 1911, corresponds to the RHS sample. The third cohort, those born between 1917 and 1921, are the so-called

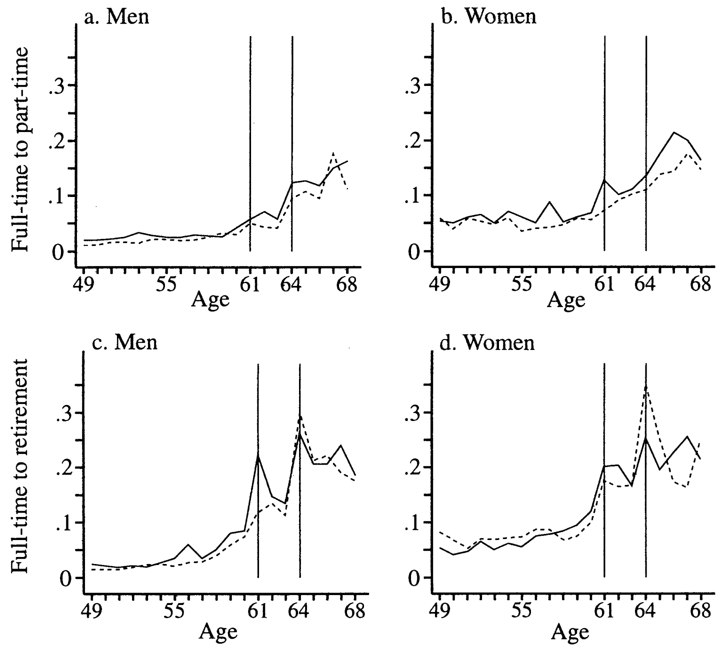

FIG. 4.—Transition rates by age and period

notch babies. For this cohort, the 1977 amendments to the Social Security Act created a substantial, unanticipated reduction in benefits with respect to those born before 1917.

The RHS is by far the most frequently studied data source on retirement. Yet for the RHS cohort, our matched CPS data show only minimal exit rates at age 61, followed by a huge spike at age 64. Over time, as we move to progressively more recent cohorts, we see the age 61 spike emerge. What at the outset of our data was a single-peak hazard, soon becomes doubly peaked. The latest cohort that we see reaching age 61 has a spike at that age that is just as sharp as the age-64 spike for the previous cohort. We think this is a vivid illustration that vis-à-vis current behavior the RHS is simply outdated. The lessons learned from it may no longer apply.

C. Decomposition of Observed Changes in Participation

We now bring trends in transitions and participation together by asking how observed changes in transitions compare to observed changes in participation.

In table 2 we present a decomposition of cross-sectional changes in full-time participation between 1968–71 (period 1) and 1987–90 (period 2) based on the identity

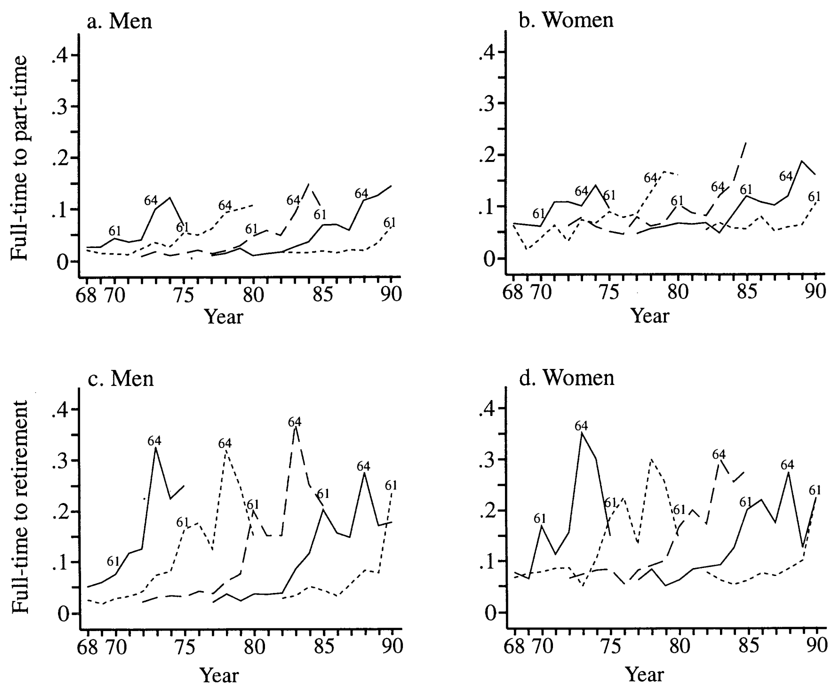

FIG. 5.—Transition rates by year and cohort

$$\pi_1 - \pi_2 = (\pi_1 - \pi_1^*) + (\pi_1^* - \pi_{12}^*) + (\pi_{12}^* - \pi_2^*) + (\pi_2^* - \pi_2),$$

where $\pi_t = \{\pi(49 + h, t)\}$ denotes the cross-section age-participation profile in period $t = 1, 2$, estimated from the unmatched CPS, $\pi_t^* = \{\pi^*(49 + h, t)\}$ denotes the synthetic age-participation profile based on the cross-section profile and age 49 participation in period t,[10] and

[10] From (1), the vector of state probabilities h years forward is given by $\pi(a + h, t + h) = \Lambda_h(a, t)^T \pi(a, t)$, where

$$\Lambda_h(a, t) = \prod_{k=0}^{h-1} \Lambda(a + k, t + k).$$

We call $\{\pi(a + h, t + h)\}$ and $\{\Lambda(a + h, t + h)\}$, respectively, the age-participation and the age-transition profiles of cohort $t - a$. In the absence of prior information, a single panel of individuals of age a at time t can identify only the age-participation and age-transition profiles of cohort $t - a$. A single unmatched CPS file at time t can identify only the cross-sectional participation profile $\{\pi(a + h, t)\}$ at time t, while a sequence of unmatched CPS files identifies all (cross-sectional or cohort) age-participation profiles. A single matched CPS file at time t identifies the cross-

$\pi_{12}^* = \{\pi_1^*(49 + h, 2)\}$ denotes the synthetic age-participation profile based on the cross-sectional transition profile in period 2 and the age 49 participation in period 1, that is, $\pi_1^*(49 + h, 2) = \Lambda_h^*(49, 2)^T \pi(49, 1)$, where

$$\Lambda_h^*(49, 2) = \prod_{k=1}^{h-1} \Lambda(49 + k, 2).$$

Thus, $\pi_1 - \pi_1^*$ is the difference between the 1968–71 cross-section profile and the synthetic one using 1968–71 transitions, $\pi_1^* - \pi_{12}^*$ is the change associated with shifts in the transition matrices between 1968–71 and 1987–90, $\pi_{12}^* - \pi_2^*$ is the cumulative effect of changes in age 49 participation valued at final period transitions, and $\pi_2^* - \pi_2$ is the difference between the 1987–90 cross-section profile and the synthetic one using 1987–90 transitions. The differences $\pi_1 - \pi_1^*$ and $\pi_2 - \pi_2^*$ may be interpreted as measures of the cohort effects at the start and the end of our data, respectively.

Table 2 shows a dramatic reduction in participation of older men, alongside what are by comparison fairly minor shifts for women. For men, the outstanding feature of the change is its concentration between ages 60 and 65. In 1968–71 the majority of men were full-time participants through age 64. By the end of the period, the majority participated only through age 61. Age 62 replaced age 65 as the median age of exit from full-time participation.

The decomposition of the change for men shows first that the initial cross-section profile π_1 cannot be supported by the initially observed transitions. By projecting the initial age 49 status to higher ages using initial transitions, we understate cross-sectional full-time participation by 6–12 percentage points in the 60–64 age range. This is one way of showing that

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sectional age-participation profiles at t and $t + 1$ and the cross-sectional age-transition profile $\{\Lambda(a + h, t)\}$ at time t, while a sequence of matched CPS files identifies all age-participation and all age-transition profiles. Given a single cross-sectional transition profile at time t, one can project participation h periods forward by defining $\pi^*(a + h, t) = \Lambda_h^*(a, t)^T \pi(a, t)$, where

$$\Lambda_h^*(a, t) = \prod_{k=0}^{h-1} \Lambda(a + k, t).$$

We call $\{\pi^*(a + h, t)\}$ the synthetic age-participation profile based on the cross-section transition profile and age a participation at time t. If the different cohorts that constitute a cross section have had different labor force histories, then the cohort, cross-sectional, and synthetic profiles will be different in general. The three reference age and age-specific transition probabilities are the same across cohorts.

Table 2
Components of Change in Full-Time Participation of Older Men and Women between 1968–71 and 1987–90

	Age						
	55	60	62	63	64	65	68
Full-time participation of men:							
1968–71	89.9	83.7	72.3	66.6	62.8	44.9	23.0
1987–90	82.2	65.0	46.0	39.2	32.3	22.8	13.0
Accounting for change:							
$\pi_1 - \pi_1^*$	1.5	5.0	4.2	5.3	6.9	3.9	3.4
$\pi_1^* - \pi_{12}^*$	5.8	13.1	22.6	21.8	22.3	17.2	7.7
$\pi_{12}^* - \pi_2^*$	1.5	.7	.4	.3	.2	.2	.0
$\pi_2^* - \pi_2$	−1.2	−.0	−.9	−.0	1.2	.9	−1.2
Total	7.6	18.7	26.3	27.4	30.6	22.1	10.0
Full-time participation of women:							
1968–71	41.5	37.7	28.6	24.7	18.9	13.6	8.0
1987–90	45.2	34.6	22.8	19.4	16.3	11.7	6.0
Accounting for change:							
$\pi_1 - \pi_1^*$	.5	2.5	−.5	−1.1	−2.0	−.3	−.4
$\pi_1^* - \pi_{12}^*$	−4.1	−2.3	3.4	6.9	5.8	3.1	3.5
$\pi_{12}^* - \pi_2^*$	−5.0	−2.1	−1.2	−.8	−.6	−.4	−.1
$\pi_2^* - \pi_2$	4.9	4.9	4.1	.3	−.7	−.5	−1.0
Total	−3.6	3.1	5.8	5.3	2.6	1.9	2.0

NOTE.—The identity for the change between the initial and the end period is

$$\pi_1 - \pi_2 = (\pi_1 - \pi_1^*) + (\pi_1^* - \pi_{12}^*) + (\pi_{12}^* - \pi_2^*) + (\pi_2^* - \pi_2),$$

where

$\pi_1 - \pi_1^*$ = difference between the 1968–71 cross-section profile and the synthetic profile using 1968–71 transitions;
$\pi_1^* - \pi_{12}^*$ = change associated with shifts in the transition matrices between 1968–71 and 1987–90;
$\pi_{12}^* - \pi_2^*$ = cumulative effect of changes in age 49 participation valued at final period transitions;
$\pi_2^* - \pi_2$ = difference between the 1987–90 cross-section profile and the synthetic one using final period transitions.

cohort participation profiles of older males were rapidly declining at the outset of our data.

The second row of the decomposition measures the effects of observed shifts in transitions between the initial and the final period. For men, the shift in transitions captures the majority of the composite change, and it clearly traces the bulge in the age 62–64 range.

The final two lines of the accounting refer to the change in participation at age 49 (a very small effect) and the difference between the final period synthetic and observed participation profiles. The negative numbers for the last difference imply that male cohort participation profiles are still declining, although not as rapidly as at the beginning of our data.

The trends for women mix increasing work outside the home with what appear to be falling ages of retirement. The net effect is small changes in full-time participation between the initial and the final periods. The patterns are interesting, however. For example, the positive numbers in the last row of the decomposition suggest that, even with the current transition rates, participation levels of women will continue to rise. The second row describes our main result from the observed changes in transitions, namely increasing full-time participation through age 60 with the largest subsequent reductions at ages corresponding to those where men also show the greatest change.

V. Labor Force Transitions: Parametric Analysis

In this section we parameterize individual transition probabilities by using a simple multinomial logit (MNL) specification. Our aim is to verify how much of our results in the previous section survive after allowing other covariates to affect the basic relationship between transition rates, age, and time. We focus on a number of personal characteristics, such as education, race, marital status, household composition, health, unemployment and self-employment status, which previous studies by several authors have shown to be important predictors of retirement.[11]

A. Model Specification

The MNL models fitted in this section are of the form

$$\ln \frac{\lambda_{ij}}{\lambda_{ii}} = f_{ij}(a, t) + x^T \beta_{ij}, \quad j \neq i, \tag{2}$$

where f_{ij} is a function of age and time only, and x is a vector of observable individual characteristics that may depend on a person's labor force status in the first survey. The population log odds ratio is modeled as the sum of two components: the first describes in a flexible way the interaction between age and time, the second is an age- and time-invariant linear function of a vector of covariates that are simply assumed to shift proportionally the odds ratio of exit from each labor force status. This MNL specification is mainly a device for smoothing the data, although it could also be interpreted as an approximation to the reduced form of some structural model of retirement behavior.

Model 2 could in principle be estimated without imposing any structure on f_{ij}. For convenience, we chose to work with a fully parametric specification, by representing f_{ij} as a time-varying cubic spline in age, with a

[11] It would take too much space to mention all these studies. We refer to Rust (1989, 1990) for a detailed survey.

single knot at age 59 and a set of dummies at age 61, 62, and 64, plus an additive business-cycle effect.[12]

Our specification allows for most of the effects discussed in the previous section, namely, nonmonotonicity and the presence of spikes in age-transition profiles, the presence of curvature in the time trends, potential differences in time trends at different ages, and the possible presence of business cycle effects.

All the covariates in the vector x are categorical. They consist of a set of indicators for race (nonblack, black), education (four levels: less than 12 years of schooling completed, or "high school dropout"; 12 years of schooling, or "high school graduate"; between 13 and 15 years, or "some college"; and 16 or more years, or "college graduate"), marital status (3 levels: married spouse present; widowed, divorced, or separated; and never married), children of less than 18 years of age living with the family, health status,[13] geographical region (Northeast, Midwest, South, and West), standard metropolitan statistical area (SMSA: in an SMSA, not in an SMSA, and missing value). For those who were not retired at the week of the first survey, we included additional indicators for self-employment and unemployed status in the week of the first survey.

We estimated model 2 separately by sex and labor force status in the first CPS year, using pooled data for the period 1968–91. At the top of table 3 we present the sample size, the number of estimated parameters, and standard measures of goodness of fit for each set of estimates. Notice that the model fits exit from full-time work better than exit from part-time work or retirement. The fit is also better for men than for women.[14]

[12] More precisely, the baseline hazard $f_{ij}(a, t)$ is of the form

$$f_{ij}(a, t) = \alpha_{0t} + \alpha_{1t}a + \alpha_2 a^2 + \alpha_3 a^3 + \alpha_4 (a - 59)_+ + \alpha_{5t}^T d + \gamma \mathrm{DGNP},$$

were u_+ denotes the positive part of u, d is the vector of age-specific dummies, and DGNP is the average second seasonal difference of log real gross national product centered at the first quarter of year t. The intercept α_{0t} is a quadratic function of time, while the coefficient α_{1t} on age and the vector of coefficients α_{5t} on the age dummies are linear functions of time.

[13] A person was classified as in bad health ("ill") if illness or disability was reported as the main reason for working less than 50 weeks the year before the survey or if illness was reported as the reason for working less than 35 hours in the week before the survey. The fact that this is not an objective but rather a self-assessed measure of health may exaggerate the effect of health on labor force transitions.

[14] An interesting question is whether the distinction between part-time participation and retirement is arbitrary in the case of exit from full-time work, and the distinction between full-time and part-time work is arbitrary in the case of exit from retirement. In our setting, testing for the pooling of states is equivalent to testing for the null hypothesis of equality of the MNL coefficients, apart from the

B. Parameter Estimates

Tables 4–7 contain the estimated MNL coefficients for all the variables included in the model. Table 3 is the analog of an ANOVA table for linear regression and reports likelihood ratio statistics for exclusion of subsets of variables. To help in interpreting the estimated coefficients on the other covariates included in the model, tables 8–10 present the baseline transition rates and the differences with respect to the baseline resulting from changing, one at a time, the value of the indicator variables in the model. Calculations are reported only for age 61 and for two different years, 1969 and 1989.

The estimates of the age and time effects, presented in tables 4–7, confirm all our findings in Section IV. In particular, there is a strong evidence of time trends in exit from full-time work and from retirement for men. The trend in exit from full-time work is concave, with the linear and quadratic term both statistically significant. In contrast, there is much less evidence of time trends in male exit rates from part-time work and in all female transition rates.

The estimates also confirm the growing importance of the spike at age 61 in exit from full-time work into retirement, especially for men, and the declining importance of the spike at age 64, especially for women.

Business-cycle effects appear to be statistically significant only in the case of male exit from full-time into part-time work and female exit from part-time work into retirement. The sign of the coefficients is negative, confirming that the hazard of leaving the labor force is highest during recessions.

Many of the findings for the other covariates agree nicely with earlier studies. Because of this, we simply highlight the role of a few covariates.

Education is a very important predictor of transition. Exit rates from full-time work and from retirement are strongly ordered by education level. People with higher education have a lower probability of leaving full-time work and a higher probability of entering the labor force from retirement. Men with higher education also have higher odds of leaving full-time work for a part-time job than retiring.

Race appears to be an important predictor in the case of exit from full-time work. Both hazards are higher for blacks than for nonblacks. Further, black women have much higher odds of leaving full-time work for a part-time job than black men, who have instead higher chances of simply retiring. We find less evidence of race effects in the other transitions, but here again, the estimated coefficients seem to indicate that black women have a greater attachment to the labor force than black men.

intercepts. We find that the null hypotheses is rejected at any conventional level of significance.

Table 3
Multinomial Logit Estimates: Likelihood Ratio Statistics for Exclusion of Subsets of Variables

	Exit from FT		Exit from PT		Exit from RT	
	Men	Women	Men	Women	Men	Women
Base model:						
No. of observations	65,803	35,619	4,746	11,353	25,546	61,714
No. parameters	56	56	56	56	52	52
$-\ln L$	21,128.2	18,431.9	4,607.3	10,685.6	7,148.1	14,321.3
$-\ln L'$	24,757.2	19,970.0	5,027.8	11,023.0	7,752.8	15,099.1
$2\Delta\ln L$	7,258.0**	3,076.2**	841.0**	674.8**	1,209.4**	1,555.6**
Pseudo-R^2	.147	.077	.084	.031	.078	.052
Likelihood ratio statistics:						
No age effects (22 restrictions)	4,611.6**	1,726.7**	577.1**	290.8**	847.3**	1,142.0**
No age dummies (12 restrictions)	271.8**	96.3**	19.4*	22.7**	15.2	17.8
No time effects (14 restrictions)	205.8**	46.9**	23.2*	40.4**	64.7**	37.4**
No age-time interactions (8 restrictions)	31.6**	19.4**	8.0	15.6**	6.3	25.3**
Only age and time effects (30 restrictions for FT and PT, 26 for RT)	2.089.8**	1,436.2**	231.9**	368.7**	514.5**	394.7**
No education effects (6 restrictions)	204.8**	168.1**	16.1**	13.1**	20.1**	95.4**
No family effects (6 restrictions)	40.1**	253.9**	22.6**	62.0**	39.7**	146.5**
No geographic effects (10 restrictions)	66.2**	34.7**	15.8	55.8**	52.4**	78.6**

NOTE.—FT = full-time work; PT = part-time work; RT = retirement; L and L' are, respectively, the maximized log likelihoods for the full model and the model with an intercept only, $2\Delta\ln L = 2\ln(L - L')$.
* = Significant at the 10% level.
** = Significant at the 5% level.

Table 4
Multinomial Logit Estimates: Age and Time Effects for Men, 1968–91

	FT to PT	FT to RT	PT to FT	PT to RT	RT to FT	RT to PT
Constant	−5.412**	−4.515**	.789**	−.915**	−.688**	−3.114**
	(.171)	(.132)	(.304)	(.358)	(.231)	(.382)
Time	.046**	.078**	.007	.018	−.093**	−.035
	(.015)	(.011)	(.026)	(.027)	(.021)	(.027)
Time^2	−.000	−.002**	−.001	−.001	.003**	.002**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Age	.201**	.104*	−.093	−.044	−.020	.189
	(.068)	(.053)	(.116)	(.143)	(.090)	(.153)
Age^2	−.032**	−.007	.008	.000	−.012	−.033
	(.011)	(.009)	(.019)	(.023)	(.015)	(.024)
Age^3	.002**	.000**	−.000	.000	−.000	−.002
	(.000)	(.000)	(.000)	(.001)	(.000)	(.001)
Agespline	−.005**	−.004**	.003	−.000	−.000	−.003*
	(.001)	(.000)	(.002)	(.002)	(.002)	(.002)
Age × time	−.001	−.000	−.000	.000	−.000	−.002**
	(.000)	(.000)	(.001)	(.001)	(.001)	(.001)
Age61	.285	.279**	.270	.089	−.413	.168
	(.188)	(.125)	(.394)	(.391)	(.452)	(.376)
Age62	−.119	.246*	−.235	.134	−.190	.063
	(.206)	(.131)	(.395)	(.346)	(.343)	(.318)
Age64	.316*	.991**	.552	.656**	−.781	−.238
	(.192)	(.124)	(.366)	(.295)	(.475)	(.305)
Age61 × time	.017	.038**	−.030	.014	−.033	.003
	(.013)	(.009)	(.027)	(.025)	(.031)	(.025)
Age62 × time	.028*	.004	.022	.016	.000	−.019
	(.014)	(.010)	(.027)	(.023)	(.024)	(.022)
Age64 × time	.010	−.017*	−.050*	−.025	.025	.011
	(.014)	(.009)	(.027)	(.020)	(.031)	(.020)
Gross national product growth	−.024**	−.000	−.004	−.010	.008	.013
	(.010)	(.007)	(.017)	(.016)	(.015)	(.016)

NOTE.—Asymptotic standard errors are in parentheses. FT = full-time work; PT = part-time work; RT = retirement.
* Significant at the 10% level.
** Significant at the 5% level.

Interestingly, the effect of family characteristics differs markedly between men and women. Men who are not married have a higher probability of leaving the labor force and a lower probability of exiting retirement. On the contrary, women who are not married have a significantly lower probability of leaving the labor force, and a higher probability of exiting retirement. Further, the presence of children in the family appears to significantly reduce the probability of leaving the labor force for men, while having only small effects for women. Overall, neglecting family characteristics leads to a major loss of fit in the case of women, especially in the case of exit from full-time work.

Geographical effects appear to be of some importance in the case of transitions from and into part-time work. In particular, living in the Northeast and in an SMSA significantly lowers exit rates from full-time

Table 5
Multinomial Logit Estimates: Age and Time Effects for Women, 1968–91

	FT to PT	FT to RT	PT to FT	PT to RT	RT to FT	RT to PT
Constant	−3.198**	−2.877**	−.807**	−1.121**	−2.969**	−3.051**
	(.123)	(.116)	(.139)	(.141)	(.138)	(.130)
Time	.028**	.003	−.027*	−.027*	−.033**	.014
	(.013)	(.012)	(.015)	(.014)	(.015)	(.014)
Time^2	−.000	−.000	.001**	.000	.002**	.000
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Age	−.077	−.054	−.169**	.008	−.110*	−.025
	(.049)	(.049)	(.055)	(.059)	(.058)	(.054)
Age^2	.018**	−.013	−.030**	.002	−.015	−.008
	(.009)	(.009)	(.010)	(.010)	(.011)	(.010)
Age^3	.001**	.001**	−.002**	.000	−.001**	−.000
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Age spline	−.003**	−.004**	.004**	−.000	.003**	−.002*
	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)
Age × time	.000	.001**	.000	.000	−.002**	−.002**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Age61	.109	.367**	.198	.056	−.009	.011
	(.207)	(.162)	(.277)	(.240)	(.302)	(.248)
Age62	.214	.120	.484*	.065	.385	.165
	(.212)	(.176)	(.294)	(.251)	(.291)	(.248)
Age64	.370*	1.043**	−.073	−.262	−.578	−.501
	(.254)	(.185)	(.352)	(.258)	(.448)	(.307)
Age61 × time	.020	.016	−.006	.022	−.005	.009
	(.014)	(.011)	(.019)	(.017)	(.022)	(.017)
Age62 × time	−.012	.017	−.043**	.012	−.039*	−.033*
	(.015)	(.012)	(.021)	(.017)	(.023)	(.019)
Age64 × time	−.010	−.029**	.000	.035**	.034	.017
	(.018)	(.013)	(.024)	(.018)	(.030)	(.021)
Gross national product growth	−.010	.003	−.015	−.026**	−.000	.001
	(.009)	(.009)	(.011)	(.010)	(.012)	(.010)

NOTE.—Asymptotic standard errors are in parentheses. FT = full-time work; PT = part-time work; RT = retirement.
* Significant at the 10% level.
** Significant at the 5% level.

into part-time work, while living in the South and in the West significantly reduces retention rates in part-time work for women.

Unemployment status is one of the most important single predictors of transition, along with age, education, and health status. For both men and women, unemployment significantly increases the probability of leaving the labor force. In particular, retention rates in full-time and part-time work drop very sharply for unemployed workers, while exit rates into retirement increase dramatically even at younger ages. This raises the question whether the distinction between unemployment and being out of the labor force is meaningful for older workers.

To check stability over time of the estimated MNL coefficients, we also fitted the model

Table 6
Multinomial Logit Estimates: Other covariates for Men, 1968–91

	FT to PT	FT to RT	PT to FT	PT to RT	RT to FT	RT to PT
High school dropout	.256**	.176**	.027	.179**	.011	−.054
	(.058)	(.040)	(.094)	(.089)	(.083)	(.086)
Some college	.047	−.163**	−.317**	.221*	−.057	.212*
	(.082)	(.060)	(.139)	(.124)	(.129)	(.124)
College graduate	−.114	−.543**	.062	.213*	.265**	.350**
	(.073)	(.057)	(.123)	(.118)	(.120)	(.118)
Black	.305**	.219**	−.199	.272**	.067	.079**
	(.090)	(.063)	(.147)	(.131)	(.125)	(.131)
Divorced/ separated	.141*	.100*	−.062	−.007	−.177*	−.219*
	(.083)	(.059)	(.123)	(.116)	(.105)	(.115)
Never married	.508**	.247**	−.613**	−.116	−.744**	−.204**
	(.105)	(.083)	(.166)	(.153)	(.160)	(.150)
Children	−.093	−.120**	.228**	−.032	.181**	.037
	(.061)	(.045)	(.100)	(.104)	(.087)	(.106)
Ill	.686**	.890**	−.405**	.350**	−1.335**	−1.038**
	(.090)	(.061)	(.126)	(.110)	(.088)	(.101)
Midwest	.116*	−.032	.092	.025	.024	.104**
	(.069)	(.047)	(.116)	(.107)	(.102)	(.106)
South	.311**	.053	.128	.133	−.217**	.261**
	(.068)	(.047)	(.112)	(.104)	(.101)	(.100)
West	.170**	.069	−.008	.052	.021	−.048
	(.076)	(.052)	(.124)	(.155)	(.106)	(.115)
Not in a standard metropolitan statistical area	.256**	−.034	−.035	.108	.370**	.197**
	(.054)	(.040)	(.086)	(.081)	(.078)	(.079)
Self-employed	.718**	−.467**	.345**	.143*	...	...
	(.051)	(.048)	(.081)	(.078)	...	...
Unemployed	1.564**	1.886**	.440**	1.678**	...	...
	(.092)	(.060)	(.219)	(.174)	...	...

NOTE.—Asymptotic standard errors are in parentheses. FT = full-time work; PT = part-time work; RT = retirement.
* Significant at the 10% level.
** Significant at the 5% level.

$$\ln \frac{\lambda_{ij}}{\lambda_{ii}} = f_{ij}(a) + x^T\beta_{ij}, \quad j \neq i,$$

separately by CPS year and for both men and women. The function $f_{ij}(a)$ is a cubic spline in age, with a single knot at age 59 and a set of dummies at age 61, 62, and 64.

For simplicity we report only the results for exit from full-time work into retirement and for those coefficients that show clear evidence of trends. Figure 6 presents the estimates by single CPS year with two standard error bands. To enhance the understanding of trends, we again smooth the estimates using Cleveland's (1979) loess. The null hypothesis of no effect of a variable is always rejected if the horizontal line at zero never pierces the

Table 7
Multinomial Logit Estimates: Other Covariates for Women, 1968–91

	FT to PT	FT to RT	PT to FT	PT to RT	RT to FT	RT to PT
High school dropout	.308**	.348**	−.069	.150**	−.133**	−.341**
	(.050)	(.046)	(.063)	(.057)	(.064)	(.056)
Some college	−.037	−.166**	−.052	.003	.168*	.172**
	(.066)	(.065)	(.077)	(.076)	(.090)	(.075)
College graduate	−.225**	−.284**	.027	−.047	.248**	.307**
	(.070)	(.067)	(.086)	(.088)	(.104)	(.085)
Black	.330**	.150**	−.108	.007	.167	.343**
	(.070)	(.069)	(.095)	(.086)	(.100)	(.087)
Divorced/ separated	−.425**	−.531**	.270**	−.299**	.718**	.428**
	(.049)	(.046)	(.064)	(.062)	(.070)	(.062)
Never married	−.695**	−.580**	.367**	−.177	.306*	.093
	(.102)	(.086)	(.148)	(.144)	(.169)	(.149)
Children	.125**	−.022	.043	.061	.039	.034
	(.054)	(.053)	(.062)	(.062)	(.066)	(.060)
Ill	.155*	.574**	.035	.227**	−.617**	−.685**
	(.088)	(.071)	(.106)	(.095)	(.096)	(.090)
Midwest	.073	.008	.019	−.033	.007	.183**
	(.060)	(.054)	(.069)	(.066)	(.081)	(.068)
South	−.000	−.050	.249**	.241**	.061	−.028
	(.059)	(.054)	(.074)	(.070)	(.078)	(.069)
West	.242**	.074	.279**	.268**	.259**	.123**
	(.064)	(.060)	(.077)	(.074)	(.084)	(.074)
Not in a standard metropolitan statistical area	.123**	−.049	.087	.140**	.262**	.271**
	(.048)	(.046)	(.058)	(.055)	(.062)	(.054)
Self-employed	.615**	.400**	.187**	.327**	...	...
	(.065)	(.068)	(.075)	(.069)	...	...
Unemployed	1.224**	2.174**	.546**	1.849**	...	...
	(.104)	(.076)	(.183)	(.133)	...	...

NOTE.—Asymptotic standard errors are in parentheses. FT = full-time work; PT = part-time work; RT = retirement.
* Significant at the 10% level.
** Significant at the 5% level.

2 standard error band. The other horizontal line is the weighted average of the estimated coefficients, with weights equal to the reciprocal of the estimated variance of the coefficients in each year. Because observations in different matched CPS are independent, this is the minimum distance estimate under the assumption of time homogeneity.

The figure shows again the rising importance of the age 62 spike and the declining importance of the one at age 64. It also shows a trend toward increasing exit into retirement for black men. Again for men, the difference between college and high school graduates (the baseline education level) has widened, while the difference between high school dropouts and high school graduates has narrowed. Finally, we see indications of a weakening of the effect of marital status on female transitions.

Table 8
Estimated Percentage Transition Rates from Full-Time at Age 61

	Men			Women		
	FT/FT	FT/PT	FT/RT	FT/FT	FT/PT	FT/RT
Year 1969:						
Baseline	89.5	2.1	8.4	78.0	6.5	15.5
Differences for change to:						
High school dropout	−1.9	.5	1.4	−6.3	1.6	4.7
Some college	1.1	.1	−1.2	2.1	−.0	−2.0
College graduate	3.5	−.2	−3.3	4.2	−1.0	−3.2
Black	−2.4	.7	1.8	−3.8	2.1	1.6
Divorced/separated	−1.1	.3	.8	7.4	−1.9	−5.5
Never married	−3.2	1.3	2.0	8.7	−2.9	−5.8
With children	1.0	−.2	−.9	−.4	.8	−.4
Ill	−11.1	1.5	9.5	−9.0	.2	8.8
Midwest	.0	.3	−.3	−.5	.5	.0
South	−1.1	.7	.4	.6	.0	−.6
West	−.9	.4	.5	−2.3	1.5	.7
Not in a standard metropolitan statistical area	−.3	.6	−.3	−.0	.8	−.8
Self-employed	.8	2.2	−3.1	−9.1	4.1	4.9
Unemployed	−31.7	4.4	27.3	−45.0	2.9	42.1
Year 1989:						
Baseline	72.5	3.5	24.0	69.7	12.7	17.6
Differences for change to:						
High school dropout	−3.9	.8	3.1	−7.4	2.7	4.7
Some college	2.6	.3	−2.9	2.3	−.0	−2.2
College graduate	8.4	−.0	−8.4	5.2	−1.8	−3.4
Black	−4.8	.9	3.9	−5.1	3.7	1.4
Divorced/separated	−2.1	.4	1.7	9.2	−3.3	−5.9
Never married	−6.0	1.8	4.2	11.4	−5.3	−6.1
With children	2.3	−.2	−2.1	−.9	1.5	−.6
Ill	−19.9	1.5	18.4	−9.5	.1	9.4
Midwest	.2	.4	−.7	−.8	.8	−.0
South	−1.8	1.1	.7	.6	.1	−.7
West	−1.7	.5	1.1	−3.2	2.7	.5
Not in a standard metropolitan statistical area	−.1	1.0	−.9	−.6	1.5	−1.0
Self-employed	4.1	4.0	−8.1	−11.3	7.0	4.4
Unemployed	−43.2	3.2	40.0	−43.7	3.4	40.2

NOTE.—Transition rates are computed from the estimated coefficients of the three-state multinomial logit model.

VI. Trends in Labor Force Participation

In this section, we look at trends in labor force participation rates of men and women across all ages. The data are averages of individual employment rates, defined as weeks worked out of 52, and are computed from the 1966–90 unmatched March CPS files.[15]

[15] Weeks worked for those who usually work part-time (less than 36 hours per

Table 9
Estimated Percentage Transition Rates from Part-Time at Age 61

	Men			Women		
	PT/FT	PT/PT	PT/RT	PT/FT	PT/PT	PT/RT
Year 1969:						
Baseline	35.8	47.3	16.9	16.9	59.6	23.5
Differences for change to:						
High school dropout	−.5	−1.9	2.5	−1.5	−1.6	3.1
Some college	−8.2	2.8	5.4	−.7	.5	.2
College graduate	.0	−2.8	2.8	.6	.4	−.9
Black	−6.1	.6	5.5	−1.5	1.0	.5
Divorced/separated	−1.4	1.1	.3	5.4	.5	−5.9
Never married	−12.1	10.6	1.5	6.6	−2.1	−4.5
With children	5.6	−3.8	−1.8	.4	−1.3	.9
Ill	−10.7	2.4	8.3	−.5	−3.7	4.2
Midwest	2.0	−1.8	−.2	.4	.3	−.7
South	2.1	−3.2	1.1	2.6	−6.0	3.4
West	−.5	−.3	.8	2.9	−6.7	3.8
Not in a standard metropolitan statistical area	−1.5	−.3	1.8	.7	−2.9	2.2
Self-employed	7.3	−7.0	−.3	1.2	−6.7	5.5
Unemployed	−7.0	−22.8	29.9	−4.6	−34.6	39.2
Year 1989:						
Baseline	15.7	54.2	30.0	14.9	55.6	29.5
Differences for change to:						
High school dropout	−.5	−3.2	3.7	−1.5	−2.0	3.5
Some college	−4.6	−1.7	6.3	−.6	.5	.2
College graduate	−.3	−4.1	4.3	.6	.5	−1.1
Black	−3.6	−3.3	7.0	−1.3	.8	.6
Divorced/separated	−.8	.6	.1	5.2	1.7	−7.0
Never married	−6.2	6.4	−.2	6.2	−1.0	−5.2
With children	3.4	−1.6	−1.8	.3	−1.4	1.1
Ill	−6.0	−3.7	9.7	−.6	−4.1	4.8
Midwest	1.1	−1.2	.0	.4	.4	−.8
South	1.1	−3.3	2.2	2.1	−6.1	3.9
West	−.4	−.8	1.1	2.4	−6.8	4.4
Not in a standard metropolitan statistical area	−1.0	−1.5	2.5	.5	−3.0	2.6
Self-employed	4.3	−5.4	1.2	.8	−7.0	6.2
Unemployed	−5.5	−31.6	37.1	−5.3	−34.9	40.2

NOTE.—Transition rates are computed from the estimated coefficients of the three-state multinomial logit model.

Figures 7*a* and 7*b* trace participation through time for men and women. Six age intervals are reported, for years 25–39, 40–54, 55–61, 62–64, 65–67, and 68–69. Within each age interval the rate reported is the simple average of the age-specific rates. The unconnected series of points labeled "All" refers to the simple average of the specific age 25–69 rates.

week) are counted as half-weeks. Within single years of age, participation rates are weighted using CPS person weights. Rates for 5-year age brackets are unweighted averages of single age rates.

Table 10
Estimated Percentage Transition Rates from Retirement at Age 61

	Men			Women		
	RT/FT	RT/PT	RT/RT	RT/FT	RT/PT	RT/RT
Year 1969:						
Baseline	5.1	5.8	89.1	1.6	2.5	95.9
Differences for change to:						
High school dropout	.0	−.3	.2	−.2	−.7	.9
Some college	−.3	1.3	−1.0	.3	.4	−.7
College graduate	1.3	2.1	−3.4	.4	.8	−1.3
Black	.3	.4	−.7	.3	1.0	−1.2
Divorced/separated	−.7	−1.1	1.8	1.6	1.2	−2.8
Never married	−2.7	−.9	3.5	.6	.2	−.8
With children	.9	.1	−1.1	.0	.0	−.1
Ill	−3.6	−3.6	7.2	−.7	−1.2	1.9
Midwest	.0	.6	−.7	.0	.5	−.5
South	−1.0	1.7	−.7	.1	−.0	−.0
West	.1	−.3	.1	.5	.3	−.8
Not in a standard metropolitan statistical area	2.0	1.0	−3.1	.5	.7	−1.2
Year 1989:						
Baseline	2.6	2.5	94.9	1.1	2.8	96.1
Differences for change to:						
High school dropout	.0	−.1	.1	−.1	−.8	.9
Some college	−.2	.6	−.4	.2	.5	−.7
College graduate	.7	1.0	−1.7	.3	1.0	−1.3
Black	.2	.2	−.4	.2	1.1	−1.3
Divorced/separated	−.4	−.5	.9	1.1	1.4	−2.5
Never married	−1.4	−.4	1.8	.4	.3	−.6
With children	.5	.0	−.6	.0	.0	−.1
Ill	−1.9	−1.6	3.5	−.5	−1.4	1.9
Midwest	.0	.3	−.3	.0	.5	−.6
South	−.5	.7	−.2	.0	−.0	.0
West	.0	−.1	.0	.3	.3	−.7
Not in a standard metropolitan statistical area	1.1	.5	−1.6	.3	.8	−1.1

NOTE.—Transition rates are computed from the estimated coefficients of the three-state multinomial logit model.

The combined relationships for men of all ages trace a step function in which participation dropped sharply in 1975 and again in 1982. Except for the 62–64 age interval, where participation continues to fall, there has been an apparent leveling in participation of men after 1982.[16]

[16] Notice that, immediately prior to 1975 and 1982, observers of aggregate participation of men could have pointed to an apparent leveling that, if used for forecasts, would have missed the marked reductions to be realized subsequently.

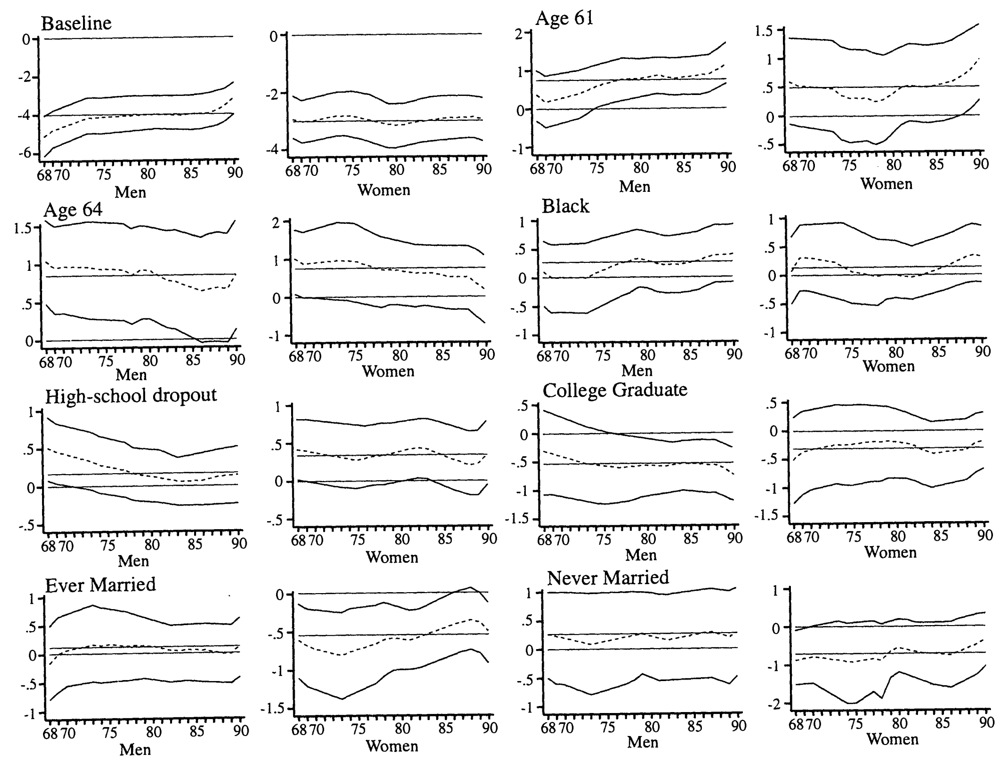

FIG. 6.—Transitions from full-time to retirement

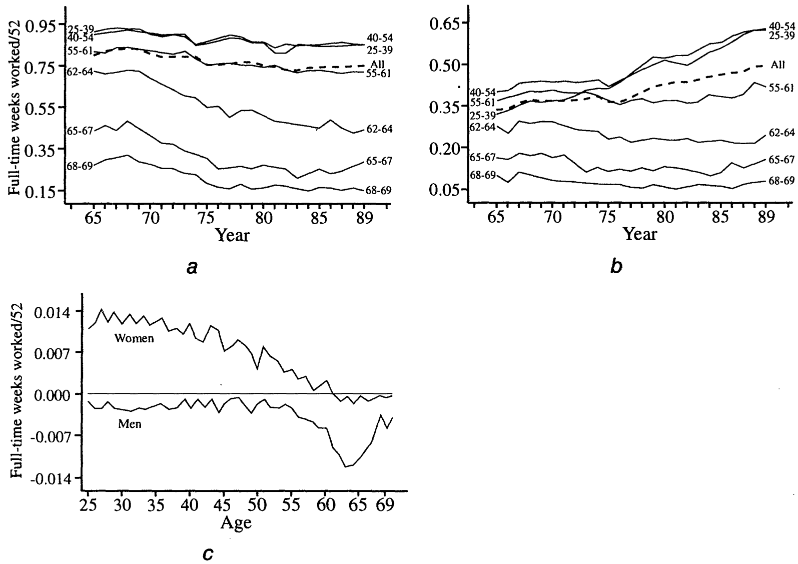

FIG. 7.—*a*, Trends in participation by age, men. "All" refers to the unweighted average of age 25–69 rates. *b*, Trends in participation by age, women. "All" refers to the unweighted average of age 25–69 rates. *c*, Patterns of changing participation by gender. Average annual change in age-specific participation, 1965–89.

The pattern for men is more or less consistent for all age groups. Prior to 1968, participation increased at all ages, but afterward the rates decline, and there are pronounced drops in the major recessions of 1975 and 1982.

Patterns for women differ by age. Participation has increased for those less than 55 years old but have not increased for older women.[17] The cyclic shocks are less clear for women than for men.

Figure 7*c* collapses the end-point data from the earlier two panels. It shows 1965–89 average annual rates of change in participation by single year of age for men and for women. Participation fell for men of all ages, and the average rate of change is roughly the same for men aged 25–54. By far the most dramatic changes have been for older men. The opposite pattern holds for women. Rates of change vary with age, and most of the

[17] To observers of participation trends the most surprising feature of this graph is the steady increase in average participation of women. Our index is the simple average of the age-specific participation rates with part-time employment discounted to full-time equivalence. As such, our index is not sensitive to changes in the age distribution of employment, but it is sensitive to shifts between part- and full-time work. The participation rates reported by the Bureau of Labor Statistics (BLS), which are frequently cited by other agencies, refer to a pool of individuals across several ages where each is distinguished as either in or out of the labor force. As such, the BLS statistic is sensitive to changes in age distributions but is not sensitive to changes in time worked. The BLS statistic shows a marked slowing in the rate of women's participation growth. The adjusted statistic does not.

action is at younger ages. Changes for women aged 60 and over have been trivial.

Figure 8*a* graphs the data shown in figure 7*a* with two changes. First, the participation rates are transformed to log-odds ratios. Next, the averages within each age group are subtracted from the observed levels. The log-odds transformation produces similar scale, while the subtraction forces each of the age-group series to have the same average value of zero. The idea is to remove scale and level so that similarity of the trends can be inspected.

The time paths are remarkably similar for men of different ages. Not only are cyclic shocks congruent, the scaled trends are as well.[18] Thus, the forces shaping employment for younger men do not appear to be fundamentally different from the forces determining the participation behavior of the oldest. In either case, relative to nonmarket alternatives, work is not as attractive to as many men today as it was in the late 1960s.

This fact is part of the reason for our opening statement that we are not convinced that wealth effects as they are ordinarily defined, specifically those that afford increased postretirement consumption through Social Security and private pensions, can be assigned a major role in explaining trends in labor supply behavior of older men. One could perhaps argue that, with pensions and Social Security in place, younger workers now understand that they need not work so hard when they are young to provide for their old age. In this view, the pension-wealth effects are not only not capable of being arbitraged via private savings, but the spillover reaches even the youngest workers.

Although we view such behavior as highly unlikely, the model is perhaps worth elaboration. Begin by thinking of an individual's average lifetime wage (the constant wage that has the same lifetime value as the individual's actual wage profile). Due to the ceiling on taxable wages and the highly progressive benefit structure, the fraction of Social Security wealth on total lifetime wealth is almost certainly a declining function of lifetime wage. Private pension wealth may be as well, in the sense that many of the most generous pension plans have been for government employees and blue-collar unionized workers who do not top the lifetime wage scale. Thus, if the importance of pension wealth (public and private) declines as lifetime earnings increase, and if the level of pension wealth increased significantly

[18] The major exception to this general statement is that cyclic shocks, both the downturns in recessions and the subsequent rebounds, are more pronounced for younger men. It can be shown that the commonality of falling employment in figure 8*a* is not an artifact of falling weeks worked for younger men with an increasing proportion of older men who do not work at all. Instead, in the period after 1968, the proportion of not working fell for men of all ages, as did weeks worked among those who work.

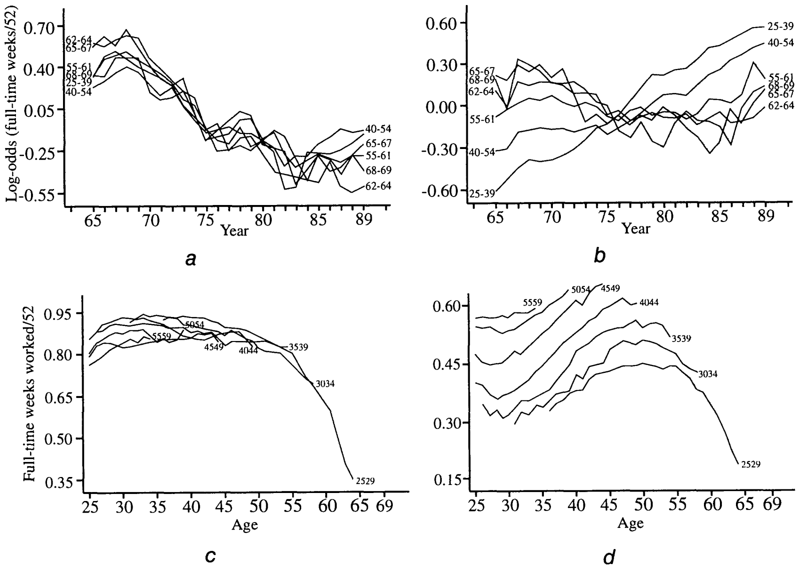

FIG. 8.—*a*, Alternative measures of trends, men (differences from age-specific average log-odds). *b*, Alternative measures of trends, women (differences from age-specific average log-odds). *c*, Cohort differences in participation, men. The labels show the range of birth years for each cohort. *d*, Cohort differences in participation, women. The labels show the range of birth years for each cohort.

during the past 2 decades, then what appears to be a wage story may be spurious, and pensions and Social Security may be the culprit after all.

Figure 8*b* is produced for comparison with figure 7*b*. It shows what we already know: women are different. Trends for younger women do not parallel those for men or for older women. If the future participation of older women is of interest, the question is, What will the participation rates of today's women who are below age 55 be in 10 years? When these women reach the ages at which today's women retire they will have had more extensive work histories than any previous cohorts.

Figures 8*c* and 8*d* complement the evidence presented so far by illustrating the behavior of male and female participation rates across cohorts.[19] In a sense, women of the younger cohorts look more and more like men in their participation behavior, although male participation rates will probably still provide an upper bound on female participation at all ages. As a consequence of the strong positive trend in young women's participation, participation rates in the 50–60 age range are rising, as successive cohorts of women with an increasing history of participation reach the older ages.

[19] Since our data cover the period 1965–89, the older cohort is observed only over the age range 36–64. For the same reason, the younger cohort is observed only over the age range 25–34.

VII. Final Remarks

Matched CPS data ought to be considered an important source of information about labor force behavior. In particular, these data make it possible to distinguish between changes in labor force transitions and changes in participation at an initial age—a distinction that we believe is important in order to understand the different trends of older men and women.

Some of the patterns that we find are not surprising. For example, the strong dependence of labor force transitions on age and other personal characteristics, such as education, race, health, family type, and composition, is well known. Other patterns are more surprising.

First, the negative trend in male participation rates appears to have flattened in the 1980s, although we face the issue of whether this is merely the result of the long expansion following the 1980–82 recession. The one exception to this is the behavior of men aged 62–64, the Social Security early retirement ages. For this age group, labor force participation rates have continued to fall, and exit rates from full-time employment have continued to rise. If credit market imperfections are the explanation for retirement at age 62, then one ought to explain why an increasing fraction of workers is liquidity constrained.

Second, as exit rates from the labor force have increased through time for older men, some of the drop points to characteristics usually associated with low wages and presumably low wealth, such as being black, having low education level, or Southern residence.

Third, aside from persistent differences in levels, the trend toward reduced labor force participation is common to men of all ages. Thus it seems that searches for explanations of changes in male participation ought to concentrate on phenomena that are common to men of all ages.

The observed negative correlation between trends in male participation and trends in Social Security benefits has often being given a causal interpretation (see, e.g., Hurd 1990). The trend toward early retirement has been also associated with changes in private pension plan rules (Ippolito 1990). Although these explanations may be part of the story, we believe that something more must be going on. An obvious candidate is increased wage dispersion. Recent work on skill-related wage differentials for various segments of the labor force (see, e.g., Murphy and Welch 1991; Juhn, Murphy, and Pierce 1993) suggests that the trends in labor force participation during the 1980s may be associated with declines in market opportunities for the less skilled workers.

A very active area of current research tries to analyze the determinants of the increase in the returns to skill. The two main explanations, increasing degree of openness of the U.S. economy and nonneutral technical changes (the two explanations are not necessarily alternative to each other), have implications for the allocation of skills in the economy, both between and within industries, and therefore may help interpret the observed differences in employment trends of older workers by education level and by industry and occupation.

An interesting question is whether changes in the industry/occupation mix have been in the direction of increasing or reducing the trend toward increased pension coverage. Answering this question may help to evaluate another proposed explanation for falling male participation rates, namely increasing coverage and generosity of private pensions. Although the CPS does not contain specific information on a firm's pension plan, it does contain information on pension availability and coverage starting with the 1979 survey. Since the March CPS contains fairly detailed information on personal and household income, evidence on the role of pensions and Social Security may also come from the analysis of trends in the structure of pre- and postretirement income.

Women's participation rates are much harder to explain. Although there is evidence of leveling in participation rates of women aged 60 or more (and, contrary to what we see for men, there is even some evidence of a rebound), the fact is that participation of these women has changed only trivially during the last 25 years. Almost all of the action in increasing participation has occurred at younger ages. Trends for younger women do not parallel those for men or for older women. When these women reach the ages at which today's women retire, they will have had more extensive work histories than any previous cohorts. Contrary to recent experience, we expect their participation rates to increase.

The observed negative correlation in the time series of male and female participation raises yet another problem. At the aggregate level, it suggests the presence of a substitution effect. At the household level, however, the evidence is of complementarity. Men who do not work are more likely to have a spouse who also does not work.

Increasing labor force participation of the younger cohorts of women is associated with changes in marital and fertility decisions. Because of this, the effects of family background variables, such as marital status, spouse's income, and number of children, on female participation probabilities and labor force transitions are also likely to change in the future. Exploring trends in these relationships may also be worthwhile.

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