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Wage and Earnings Profiles at Older Ages

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Abstract

The inverted U shape of the lifetime wage profile is frequently taken to be a stylized fact. This implies a smooth decline in wages as workers approach retirement. Instead, this paper shows that the hourly wage of the typical older worker increases slightly with age for as long as he is employed full time. It declines discretely only when he enters partial retirement, a transitional period characterized by the prevalence of part-time work, and remains mostly flat thereafter. That is, the wage path at older ages is best represented by a step function. The smoothly-declining profile often found in the literature is the result of aggregation over individuals who enter partial retirement at different ages. This conclusion is robust to controlling for self-selection into partial and full retirement. More importantly, the transition out of full time work is a choice for most workers, and the subsequent wage change is endogenously determined. While standard labor supply models would rationalize the reduction in hours worked upon partial retirement as a response to an exogenously declining wage trajectory, the evidence presented in the paper indicates instead that workers choose to trade more leisure for a lower hourly wage in a context in which a better paid, full-time job is available. In other words, wages and hours are jointly determined at older ages. These findings have important implications for the analysis of saving and labor supply decisions over the life cycle.

JEL Codes: J30, J26

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1 Introduction

The age-wage profile of a typical individual, also known as the *deterministic* or *predictable* component of wages, is a key input in empirical labor and macroeconomic models. But what does this profile look like? Labor supply studies generally assume it has an inverted U shape: wages increase in the early stages of a worker's career, flatten off in middle-age, and finally decline as he approaches retirement. Indeed, the result that wage rates decline smoothly with age for older workers is so ubiquitous in the literature that it is frequently taken to be a stylized fact.

This paper shows that there is no empirical evidence of wage declines at any age for full-time male workers. On the contrary, the hourly wage of the typical man over age 50 increases slightly with age for as long as he is employed full-time. The wage then drops discontinuously when the worker transits into partial retirement -a state characterized by the prevalence of part-time work-, and remains mostly flat thereafter. Thus, the wage path at older ages is best represented by a step-function. The smoothly-declining profile often found in the literature is a result of aggregation over workers who transit into partial retirement at different ages.

Because models of life-cycle consumption and savings often use labor earnings, rather than wages, as an input,[1] the paper also studies the age-earnings profile. This is shown to have the same step-function form as the wage profile, with a much more pronounced drop (73% vs. 32%) upon partial retirement due to the simultaneous discrete change in hours worked.

As an additional contribution, the paper analyzes the nature of the transition into partial retirement and its implications for labor supply modeling. I begin by showing that the finding on the shape of the age-wage profile is not a result of self-selection. That is, the non-decreasing wage path of full-time workers is not driven by individuals who experience negative wage shocks choosing disproportionately to (partially) retire. Moreover, the transition out of the full-time job is voluntary in the majority of cases. Altogether, the evidence points to workers who partially retire *choosing* to trade more leisure for a lower hourly wage, rather than reducing hours in response to an exogenously declining wage path that makes full-time work less attractive every year.

These results challenge the standard treatment of observed hours and wages in labor supply studies. The canonical labor supply model assumes that a worker selects the number of hours worked that maximizes utility for a given wage. The finding that workers choose to switch to a part-time job in a context in which a higher-paid, full-time job is available suggests, instead, that the typical worker faces a choice over a bundle of hours and wages. The appropriate approach to modeling labor supply at older ages must incorporate the joint decision on the two variables.

The implications of these findings are wide-ranging because key results in the labor and macro literatures depend on the correct specification of the age-wage profile. For example, the intertemporal elasticity of substitution in labor supply (IES) is an important parameter for the study of

[1] This is a simplification intended to avoid modeling the labor supply decision.

labor supply responses to business cycles and tax, private pension, or Social Security reform. The estimation of this parameter using panel data relies crucially on the distinction between movements along the life-cycle wage path and unanticipated changes that switch the profile itself. Incorrectly assuming an exogenously declining wage profile at older ages will yield biased estimates of the IES and consequently inaccurate policy predictions.[2]

The same is true for estimates of income uncertainty, which are key for the study of life-cycle consumption and savings. It is common in this literature to approximate the predictable component of earnings or wages as smoothly declining from the late 50's (see Hall and Mishkin [1982], Hubbard, Skinner, and Zeldes [1994] or Floden and Lindé [2001]), and implicitly assume that deviations from this path are the result of shocks. This approach leads to an overestimation of the labor income risk facing older individuals if wages do not decline at older ages (as is the case for full-time workers) or if declines are endogenously determined (as is the case upon partial retirement).

The wage profile is also a pivotal component of retirement models. Wages have a direct effect in the retirement decision, in that they determine the relative value of work versus leisure. If partial retirement is a choice for most workers, models that treat it as the optimal response to an exogenously declining wage profile are misspecified. Retirement transitions driven by factors such as the worker's taste for leisure or other financial incentives, like private pensions and Social Security, will be wrongly attributed to the anticipated decline in full-time wages. Expected wages also affect the retirement decision indirectly, through their effect on pension accrual. Computations of present discounted values of private pension and Social Security benefits for nonretired individuals often assume their wage follows a downward-sloping path (see Poterba, Rauh, Venti and Wise [2007]).

Finally, the shape of the wage profile is important for our understanding of the functioning of the labor market. Different theories of wage determination have different implications concerning the relationship between productivity and wages as the worker ages. The smoothly declining wage profile is often justified on the grounds that it is consistent with the human capital model, which predicts that wages of older workers will follow the declining path of productivity. The step-function profile found in this paper can only be reconciled with the human capital model in two scenarios: either productivity does not decline at older ages, or wage rigidities prevent adjustments to reflect productivity drops. On the other hand, the finding of a non-declining wage profile for full-time, career workers is consistent with alternative wage theories such as Lazear's (1979) model of deferred compensation.

The outline of the paper is as follows: Section 2 reviews the literature on the age-wage profile. Section 3 describes the data used in the empirical analysis and provides some descriptive statistics. Section 4 presents estimates of observed age-wage profiles that control for transitions into partial

[2]The Euler equation-based approach to the estimation of the IES, pioneered by MaCurdy (1981), is most sensitive to an incorrect specifications of the wage profile, as it rationalizes hours changes as responses to changes in observed wages, taken as exogenous. The life-cycle approach in Rogerson and Wallenius (2009), instead, can endogenize discontinuous changes in hours as a choice between a better-paid full-time job and a lower-paid part-time job.

retirement. Section 5 tests for bias in the observed profiles resulting from non-random selection. Section 6 discusses what proportion of partial and full retirements is voluntary, and section 7 concludes.

2 Overview

The standard labor supply model posits that participation and hours choices are a function of an exogenous wage rate (Heckman [1993]). Not surprisingly given their key role, a large literature exists on the study of wages. This literature has paid careful attention to wage uncertainty, but comparatively less focus has been placed on age-wage profiles, particularly as workers approach the end of their careers.

Labor supply models generally assume an exogenous life-cycle wage path that peaks in middle age and declines smoothly thereafter.[3] However, despite its pervasiveness in the labor and macro literature, the empirical evidence of a decreasing wage profile for older workers is plagued with problems.

The justification for the assumed shape of the lifetime wage path dates back to the human capital model (Becker [1993], Ben-Porath [1967], Mincer [1974]). The most basic version of this model predicts that investment in human capital will dominate depreciation in the early years of a worker's career, leading to increasing productivity. As he ages, the worker's incentives to invest in human capital diminish. Thus, the slope of the productivity profile flattens off with age and eventually declines. If hourly wages are a direct translation of productivity, this gives rise to an inverted U-shaped wage profile.

In the absence of depreciation, however, the productivity profile would just flatten off but never decline. Moreover, alternative theories of wage growth over the life cycle, such as models of deferred-compensation, do not predict a close association between productivity and wages or, for that matter, a declining wage profile for older workers. The shape of the lifetime wage path is, therefore, an empirical question.

Most of the empirical research on wage profiles dates back to a time when panel micro data sets were not available. These cross-sectional studies find wage declines that set in relatively early, generally while workers are still in their 40's.[4] But cross-sectional regressions characterize the

[3] Models that abstract from modeling labor supply assume that it is earnings, rather than hourly wages, that decline at older ages. For the time being, it is useful to think as labor supply as fixed, which makes the focus on wages or earnings equivalent. The role of labor supply is introduced in section 4.2, which shows that the conclusions of the paper apply to both wages and earnings.

[4] Lillard and Willis (1978), Mincer (1974) and Welch (1979) analyze earnings profiles, which may overstate the decline in wages if hours decrease too. Hurd (1971) studies wage profiles and still finds that these peak in the early 40's. For other examples see the survey by Willis (1985), where he states "Earnings functions [...] have been estimated hundreds of times [...]. Almost all of the earnings function estimates that I have seen indicate a concave log earnings-experience profile". Since the traditional earnings function does not control separately for age, a concave experience profile implies a concave age profile for men, for whom actual and potential experience tend to be very

average wage profile, which will be different from the *individual* wage profile in the presence of cohort effects or composition effects. If, for instance, individuals born in later years have higher wages, or those with higher education and better jobs retire earlier, the average wage profile will decline more rapidly than the individual wage profile.[5]

The results from more recent longitudinal studies such as Honig and Hanoch (1985b) and Johnson and Neumark (1996) suggest that cohort and composition effects are indeed responsible for part of the downward trend in cross-sectional wage profiles at older ages. But even after controlling for those they find significant wage drops in their samples of older workers, even though these are not apparent until workers are in their late 50's or early 60's.

Interestingly, both Honig and Hanoch and Johnson and Neumark find sharp wage declines around the Social Security ages. This is significant because many men reduce their hours of work in those years, possibly as a response to incentives from the Social Security rules (Aaronson and French [2004]). Indeed, part-time work is a common phenomenon at older ages. The retirement literature has shown that the majority of workers do not transit directly into inactivity after leaving their long-term, career job. Instead, they go through a transitional period of partial retirement, generally working part-time, often outside their former industry and occupation.[6] There are many reasons to expect hourly wages to decline upon partial retirement. First, part-time workers may suffer an hourly-wage penalty with respect to full-time workers in similar jobs. Second, the transition into partial retirement often involves a change of employer and job characteristics, with the consequent loss of job-specific skills. These job changes are sometimes motivated by lack of flexibility of their former employer, who may not accommodate part-time work. Other times workers actively search for less demanding jobs, which they see as part of a gradual transition into full retirement.

Put together, this evidence suggests that part of the decline in the individual wage profile is likely due to transitions out of full-time jobs. To my knowledge, only Gustman and Steinmeier (1985) have estimated wage equations controlling for partial retirement status. They find that even though the transition out of the full time job is indeed associated with a significant wage drop, the wage profile of non-retired workers is still decreasing at older ages. But their OLS estimates do not control for cohort and composition effects which, as mentioned above, have been shown to be important in this context.

In order to address the issues highlighted in this section, the first part of this paper is concerned with the estimation of the wage profile of nonretired workers. Having established that this is slightly increasing, and that wage declines at older ages are fully explained by the transition into partial

close at older ages.

[5]The individual wage profile obtained from fixed effects regressions is not affected by composition effects as long as the characteristics associated to earlier transitions into retirement are fixed (e.g. education, type of pension plan, etc.). This profile may, however, still be biased as a result of selection on unobservables. This possibility is explored in section 5.

[6]See Cahill, Giandrea and Quinn (2005), Gustman and Steinmeier (1985), Honig and Hanoch (1985a), Quinn (1997), and Ruhm (1990 and 1995) for a detailed description of the retirement process and its evolution over time.

retirement, the second part of the paper examines the nature of this transition. First, it will be shown that it is not driven by non-random selection into retirement, whereby workers who receive negative wage shocks choose disproportionately to leave their full-time job. Second, I provide evidence that the majority of transitions into partial and full retirement are voluntary.

The voluntary nature of most retirements is key to understanding the implications of the paper. The fact that most workers choose to leave their full-time job when they could have kept it is of major significance for the modeling of labor supply at older ages. It shows that individuals do not face one exogenous wage rate. Instead, hours and wages are determined jointly, as the results of a choice between a full-time job with a higher wage rate, and a part-time job with a lower one.[7]

In this sense, this paper is also related to the literature that has questioned the assumption that workers have flexibility in choosing hours, and provided evidence that wages and hours are tied together (See Moffit [1984], Altonji and Paxson [1988, 1992] and Blundell, Brewer and Francesconi [2008], among others). These papers focus on younger workers. Given the substantive evidence indicating that older workers, for the most part, want to retire gradually (see Siegenthaler and Brenner [2001] for a review of this evidence), the lack of flexibility in hours choices is likely particularly binding for them. Modeling transitions into partial retirement as a response to an exogenous decline in wages misrepresents the constraints faced by older workers. Most of them choose to leave their full-time job and forego the higher wage even when they could have kept it, for reasons that a correctly-specified labor supply model needs to capture.

3 Data

The data used in the empirical analysis come from the Health and Retirement Study (HRS), a panel dataset of adults over 50 years of age. These are complemented with cross-sectional data from the 1990 US Census 5% Public Use Micro Sample (PUMS), which covers workers of all ages, wherever full life-cycle profiles are shown.[8]

In order to work with a homogeneous sample, the analysis is restricted to men born between the years 1931 and 1941. The majority of these individuals enter the sample in the first wave, corresponding to the year 1992, and are subsequently interviewed every two years. I use the 9 survey waves that are currently available, spanning from 1992 to 2008.

There are 4,856 HRS respondents who were born during the period of interest. 1,089 of those are already fully retired by the time they become HRS respondents. As there is no information on

[7]This argument assumes that workers cannot work part-time and maintain the higher, full-time wage rate. There is ample evidence in the literature that this is indeed the case for older workers. Gustman and Steinmeier (1985) provide evidence on minimum hours constraints in career jobs from the PSID. More recently, Hurd (1996) argues that employers are often unable or unwilling to accommodate flexible hours, forcing workers to move to leave their career job if they want to reduce the number of hour worked.

[8]For a description of the HRS dataset, see HRS "RAND HRS Data, Version J" (2010). For the Census PUMS, see Ruggles et al. (2010).

wages for these individuals, they must be excluded from the analysis. It is therefore important to understand the characteristics of this group. Not surprisingly, they are slightly older when they become HRS respondents than the sample of workers (57.28 vs 55.86 years old, on average). In terms of the reasons that may have led them to retire, they are somewhat heterogeneous, and can be broadly divided in three groups: 33% of them are receiving disability insurance, 27% are in bad health but not receiving disability, and 40% of them are in good health.[9] Those receiving disability are the poorest (24k median total household wealth[10]), followed by those in bad health but not disabled (44k) and those in good health (136k). For comparison, median total household wealth for the sample of workers is 90k. These figures suggest that those in good health are likely to have had good jobs with access to private pension plans that allowed them to retire early, while those in bad health or disabled may have had to exit the work force for health reasons. In terms of the effect of excluding these workers on the interpretation of the final results, it is important to bear in mind that the number of disability insurance recipients more than doubles (up to 17%) before age 60, so this form of retirement is not excluded from the sample in the main analysis.

Workers who are self-employed when first observed in the HRS panel (762) are excluded from the final sample, as well as those with hourly wages below $3 or over $40, and those for whom only one wage observation is available. The remaining sample consists of 2,253 individuals.

The dependent variable analyzed in the following section is the (hourly) *Wage Rate*. Figure 1 shows the average wage profiles from the HRS and Census (the left panel spans the whole life cycle, while the right one covers only the years for which information from both data sources is available). The Census series has the well-known inverted U shape. Average wages increase with age for young workers, flatten off from the early 40's and decline from the early 50's. HRS observations span the period of declining hourly wages. The two series do not overlap for two reasons: first, the Census hourly wage rate needs to be recovered from yearly wage earnings, which include commissions, cash bonuses, tips, and income from secondary jobs. The HRS variable, instead, measures directly the hourly wage rate at the primary job. Second, the Census series is a cross-section, while the HRS one uses pooled observations for individuals born between 1931 to 1941 at all ages. This implies that as we move towards older ages the Census series should be more affected by cohort effects. Despite these differences, the rate of wage decline is remarkably similar in the two series.

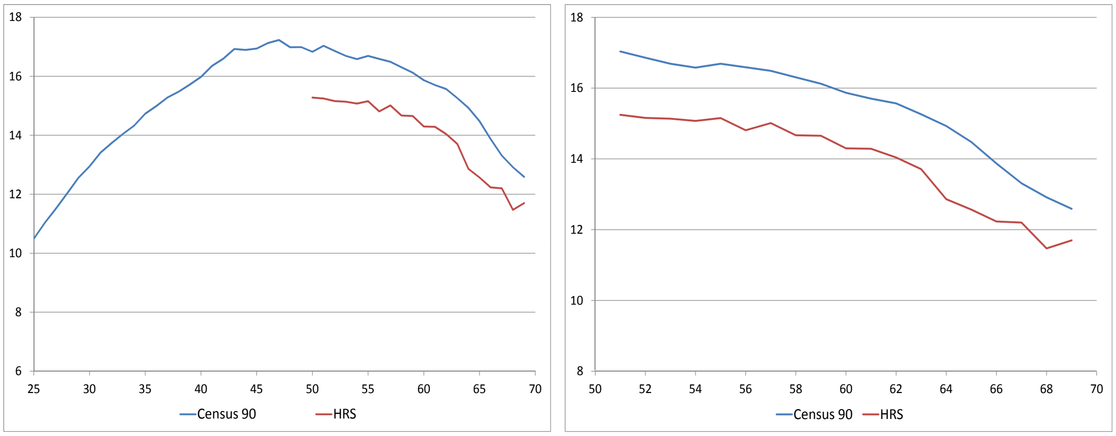
The other key variable in the empirical analysis is participation status, which is defined based on the variable *Hours of work per week at current job*. Workers are classified as working full-time if they work more than 30 hours per week, and as working part-time if they work up to 30 hours per week.[11] The left panel of figure 2 shows life cycle participation profiles from the Census. Two

[9] The proportion of disabled workers implies that 7.4% of the incoming HRS sample has retired through disability, which corresponds fairly closely to the 7% of the total workforce aged 50 to 54 receiving disability benefits in 2009 (Milligan [forthcoming]).

[10] All financial variables are expressed in 1992 $.

[11] Using different definitions of full-time work, part-time work and inactivity, such as classifying as full-time workers those who work more than 35 and considering those who work less than 2 or 5 hours per week as inactive, does not

Figure 1: Average hourly wages by age, 1990 Census and HRS.

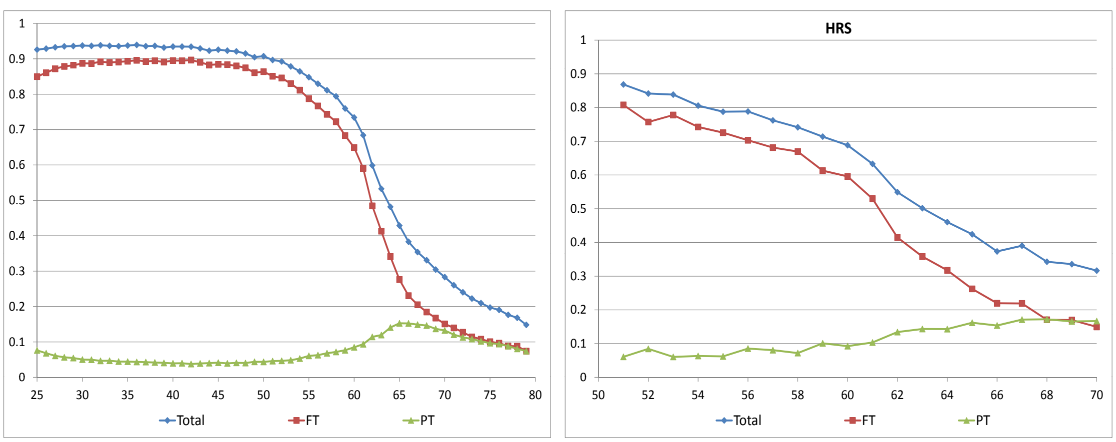


NOTE. - All figures in 1992 dollars. Hourly wages lower than $3 per hour and higher than $40 per hour excluded.

main trends are noteworthy: On the one hand, total participation rates start declining at the same time as average hourly wages. On the other hand, full-time work decreases more rapidly than overall participation, while part-time work increases from the early 50's. Thus, the full-time versus part-time composition of the work force changes radically between ages 50 and 70.

Finally, the right panel of figure 2 shows participation rates in the HRS. The changing proportions of full-time and part-time workers in the total workforce are apparent. The only difference with the Census is that full-time work declines slightly more rapidly relatively to part-time work in the HRS, reflecting a secular trend towards partial retirement (Ruhm [1995]).

Figure 2: Total, full-time and part-time participation by age, 1990 Census and HRS.



A worker's participation status is used to define partial retirement. An individual becomes

affect the results.

partially-retired whenever he is first observed working part-time or he first becomes self-employed. The decision to keep in the sample full-time workers who only become self-employed when they approach retirement follows substantive evidence that self-employment is an important route into partial retirement in the US (Maestas and Zissimopoulos [2010]). Dropping the self-employed from the sample does not affect the final results substantively, and indeed makes the effect of partial retirement on wages slightly stronger.

Partial retirement is an absorbing state, that is, partially retired individuals who go back into full-time work do not recover their "non-retired" status, as these so-called "reverse retirements" or "unretirements" are unlikely to imply a return to a career job.

4 Observed Profiles

4.1 Age-Wage Profiles

Denote by w_{it} the log hourly wage of individual i in period t. As is customary in the literature, the following linear specification is assumed for wages:

$$w_{it} = W(Age_{it}) + X_{it}\beta + u_{it}, \tag{4.1}$$

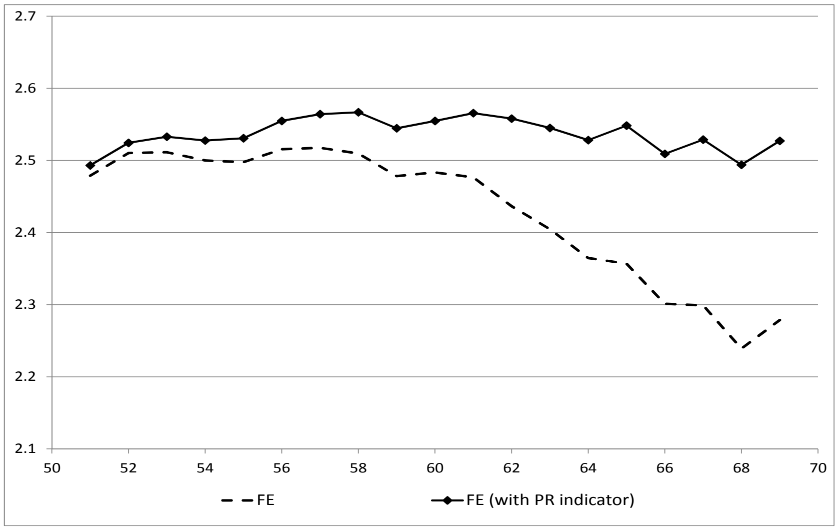
where W is some function of age, X_{it} a vector of other observable characteristics, and the error term u_{it} is potentially serially correlated and/or correlated with the observables.

Equation 4.1 can be interpreted as a standard Mincerian regression where hourly wages, rather than earnings, are used as the dependent variable to avoid confounding wage and labor supply effects. The roles of education and experience are not made explicit here to maintain the focus on age, but both are implicit in the regression: education will be captured by the fixed effect, while experience is embedded in the age effect.

I begin by estimating equation (4.1) following the conventional approach of using all available wage observations, without distinguishing between those of non-retired and partially retired workers. The second column of table 1 reports fixed effects coefficient estimates from this regression (for comparison, pooled OLS estimates are also shown on the first column). The age profile implied by these results, displayed as the dashed line in figure 3, is clearly decreasing from the late 50's onwards. As the slope changes around this age, I run two separate joint significance tests. The p-values, shown at the bottom of table 1, show no significant change in wages from ages 51 to 57, but a strongly significant decline from ages 58 to 69.

The results just presented come from a sample that includes partially retired workers. There are a number of reasons why their wage rates may be lower than those of full-time workers, including lower hourly wages for part-time work and differences in job characteristics if they move to jobs that require less effort or responsibility. As the proportion of partially-retired workers in the

Figure 3: Log-Wage Profiles for Different Specifications



sample increases tenfold from ages 55 to 65, failure to control for this composition change may be problematic.

To capture this, I re-estimate equation 4.1 adding as an additional regressor an indicator that the individual has entered partial retirement. Since this is an absorbing state, its coefficient is identified from individuals who switch from full-time work. Estimates from this regression, reported in the third column of table 1, show that entrance into partial-retirement is associated with a 32% decline in hourly wages. This result is consistent with Aaronson and French (2004), who find a 25% wage drop for individuals who change their labor supply from 40 to 20 hours per week upon becoming 62 and 65, although the effect of partial retirement captured here is more general, and not confined to the Social Security ages. More importantly, the wage profile obtained after controlling for partial retirement -shown on the solid line in figure 3- no longer declines with age: Joint significance tests show a slight increase in wages until age 57, and no significant change thereafter.

The implication of this result is that the individual age-wage profile is mostly flat or slightly increasing for workers over age 50 who remain employed full time, declining discretely only at the point of entry into partial retirement. This is illustrated in figure 4, which shows, as an example, the predicted wage profile for an individual who transit into partial retirement at age 55, on the left, and at age 62, on the right, using the coefficients from the last regression. It is clear that these step-function profiles are very different from the smoothly declining one obtained from the regression that did not control for retirement status. That profile was a result of aggregation over individuals who transit into partial retirement at different ages.

The effect of aggregation can also be illustrated using the results of separate regressions for the subsample of non-retired workers (which yields the *full-time* profile, shown on the left series in figure 5) and that of partially-retired workers (the *part-time* profile, shown on the right series in

Figure 4: Examples of predicted wage profiles.

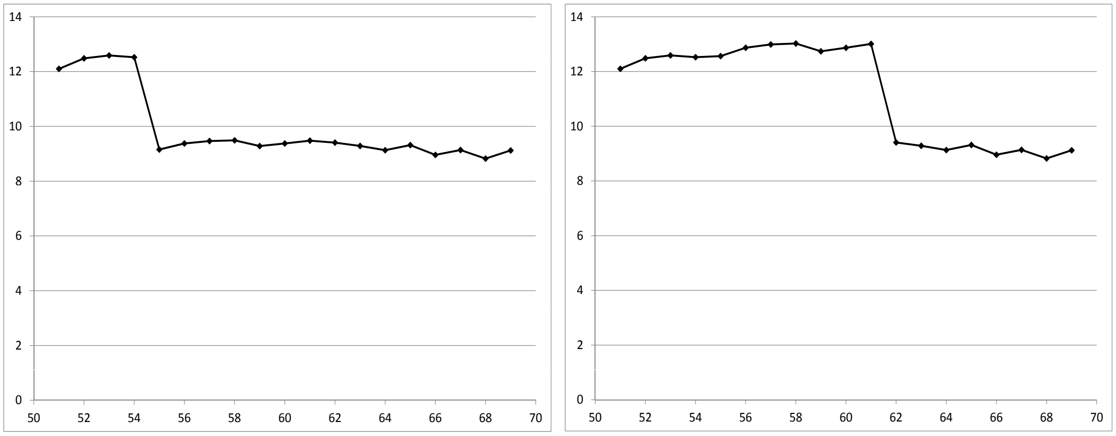
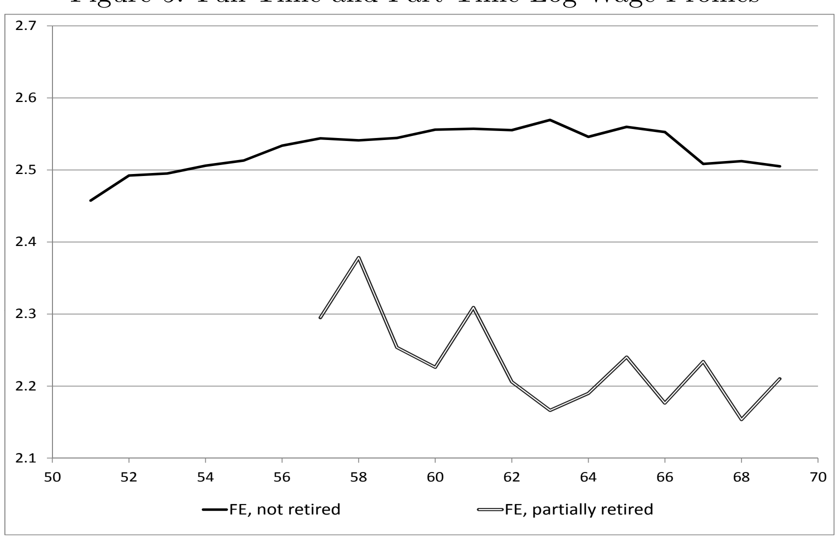


figure 5). Joint significance tests provide no evidence of wage declines for full-time workers, and slight but significant declines for partially-retired workers, which are likely explained by further job changes for those already retired. Failure to control for transitions into partial retirement in wage regressions yields a weighted average of these two profiles. As we move towards older ages, partial retirement becomes relatively more likely, and hence the weight on the part-time profile increases. As a result, the *average* wage profile decreases with age.

Figure 5: Full-Time and Part-Time Log-Wage Profiles



4.2 Age-Earnings Profiles

Life-cycle models of consumption and savings often avoid modeling the labor supply decision explicitly and take the labor earnings process, rather than hourly wages, as exogenous. They regularly assume that labor earnings of older workers decline smoothly with age. But as earnings are partly determined by the hourly wage rate, the results in the previous section imply that they must also drop discretely upon entrance into partial-retirement. And since, by definition, this transition is associated with a reduction in hours, the discrete drop in earnings is likely even more pronounced than that in wages. The results presented above regarding the shape of the wage profile are not enough, though, to infer the shape of the earnings profile of full-time workers, which will depend on whether the adjustment in hours worked begins before the transition into partial retirement. This section investigates this point, and characterizes the shape of the labor earnings profile of older workers.

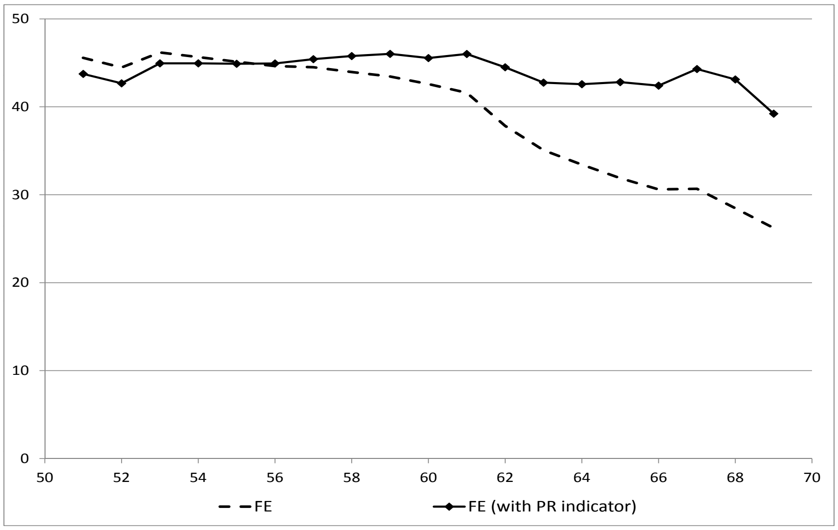
The biennial nature of the HRS survey results in a mismatch between the labor supply and labor earnings variables. At each interview, individuals are asked about their current work status and their earnings for the preceding calendar year. As a result, information on the number of hours worked is available for the interview years only. Earnings information, meanwhile, is available for the intervening years, during which it is not possible to determine which individuals are partially retired. Therefore I do not use reported labor earnings in the analysis, and instead construct a contemporaneous measure of weekly earnings combining information on hourly wages and hours of work per week at the current job. Results for wages have already been presented in the previous section. Here I first analyze the hours profile, and then discuss the implications for the earnings profile.

To study the profile of weekly hours I estimate equations of the same form as 4.1 using log hours as the dependent variable. The regression coefficients are shown in table 2, and the implied age profiles (in levels) are plotted in figure 6.

As before, I start by estimating the hours profile using all available observations, without controls for retirement status. Fixed-effects estimates are presented on the second column of table 2 (OLS estimates are shown in column 1 for comparison). The coefficients from this regression show a steeply declining hours profile that drops by more than 42% (in levels) from ages 51 to 69. There are very significant drops in hours (9.5 and 7.6%, respectively) at ages 62 and 63, confirming workers' tendency to reduce hours at the Social Security's early retirement age, as shown in Aaronson and French (2004). The average number of hours worked diminishes very slightly (by 2.3%) between ages 51 and 57, but the age dummies for these years are not jointly significant at 5%. The bulk of the decline in hours takes place between ages 58 and 69. Hours go down by 40% in this period, and the corresponding age dummies are strongly jointly significant.

To understand what proportion of the downward trend in hours is due to transitions into partial retirement, as opposed to adjustments within the full-time job, I add to the fixed effects regression

Figure 6: Hours Profiles for Different Specifications

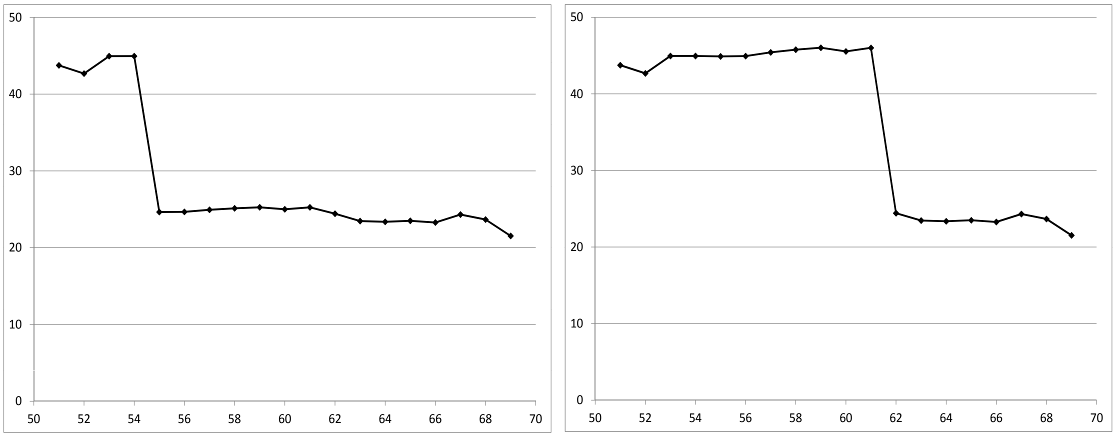


an indicator that the individual is partially retired. As shown on the third column of table 2, the coefficient on this variable is strongly significant, and implies a 60% drop in hours upon exit from the full-time job. Once we control for the transition into partial retirement, the age dummies around the Social Security ages are no longer significant and their coefficients shrink in magnitude, indicating that they were capturing outflows from full-time jobs in the previous regression. The resulting age profile is now slightly upward-sloping from ages 51 to 57, and slightly downward-sloping thereafter, providing some evidence of hours adjustments in anticipation of retirement. The overall decline in hours from age 51 to age 69 is reduced to 11%, and most of it results from the large (but not significant) negative coefficient on the dummy for age 69. In other words, 76% of the change in hours between ages 51 and 69 (and 96% between 51 and 68) is explained by the transition into partial retirement.

Figure 7 shows predicted hours profiles for a hypothetical individual who transits into partial retirement at age 55 (left) and one who does so at age 62 (right), constructed using the coefficients from the last regression. The hours profile has the same step-function shape as the wage profile. It is essentially flat for full-time workers and drops discretely with the switch into partial retirement. Further hours adjustments in the following years are comparatively minor. As was the case for wages, the smooth hours profile often reported in the literature is a result of aggregation of non-retired and partially retired workers.

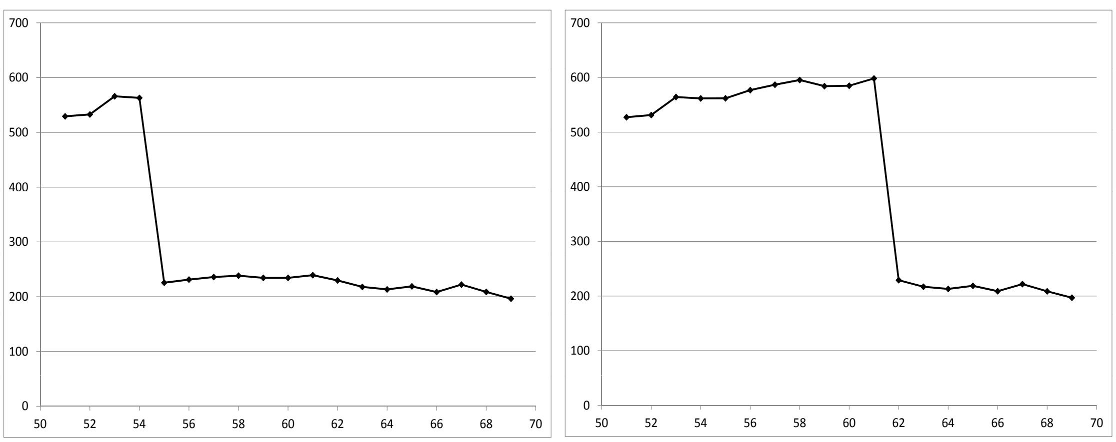
Finally, the results for wages and hours can be combined to offer some insight on the shape of the earnings profile. Figure 8 shows predicted weekly earnings profiles for an individual who transits into partial retirement at ages 55 and 62, respectively. Not surprisingly, as weekly earnings are obtained as the product of the wage and hours profiles shown above, the earnings profile is also a step function. This analysis misses some standard components of labor earnings, such as bonuses,

Figure 7: Examples of predicted hours profiles.



tips, or income from secondary jobs. It is conceivable that these may be adjusted before partial retirement. However, in view of the evidence in figure 8, it seems safe to conclude that the drop in earnings upon entrance into partial retirement accounts for the vast majority of the decline in earnings at older ages.

Figure 8: Examples of predicted earnings profiles.



5 Testing for Selection Effects

It follows from the analysis in the previous section that the observed wage profile is slightly increasing for non-retired workers. This finding appears to be inconsistent with a model where productivity declines slowly with age and wages reflect worker productivity. But the shape of the observed profile could be a result of selection on unobservables. Suppose that productivity is stochastic and follows a downward trend with age. Every year, only workers who receive positive productivity shocks keep their full-time jobs, while the rest transit into partial or full retirement. In this scenario, the observed full-time wage profile understates the decline in wages with age.

This section provides evidence to show that the flat full-time wage profile is not driven by non-random selection into retirement. The least parametric evidence comes from the fixed effects regressions in the previous section. As shown on the second column of table 1, the wage profile obtained from the regression mixing full-time and part-time observations declines at age 62 by almost 4 percent. 62 is the age at which individuals become eligible for Social Security benefits, and it has been found to be associated with a pronounced spike in the retirement hazard.[12] If wage declines are partly explained by switches from full-time work into partial retirement, we would expect to find a drop in wages at this age. Tellingly, the wage drop at 62 disappears completely once the control for partial retirement is introduced. If transitions into partial retirement at this age are largely motivated by exogenous institutional features, this result indicates that the most noticeable decline in wages at older ages is not driven by self-selection.[13]

Next, I provide more formal evidence of lack of selectivity by performing the selection test proposed by Wooldridge (1995 and 2002), which extends Heckman's (1979) two-step procedure to a panel-data context. The tradeoff for the added formality is a stronger reliance on functional form assumptions.

I consider two alternative specifications for the selection process. In the first one, I group partial retirement and full retirement together, and model the selection rule as a choice between working full time and not working full-time, with full-time wages observed only in the former case. In the second one, I consider a selection rule based on two separate indices, equal to 1 if the individual is partially and fully retired, respectively. Full-time wages are observed whenever both indices are equal to 0. This specification allows for observables to affect differentially the decisions to partially and fully retire, in line with the retirement literature that characterizes the two as separate statuses (Honig and Hanoch [1985a]).

In the single-index case, selection is determined by the following equation:

$$s_{it}^R = I[Z_{it}\beta + \bar{Z}_i\Pi + v_{it} > 0], \tag{5.1}$$

where $v_{it}|Z_i$ has a standard normal distribution, s_{it}^R is an indicator equal to 1 if the individual is retired (either partially or fully), Z_{it} is a vector of observables and $\bar{Z}_i$ a vector of within-individual averages defined as follows: $\bar{Z}_i \equiv \frac{1}{T_i}\sum_{j=1}^{T_i} Z_{it}$, with T_i equal to the number of periods that individual i is observed in the panel. Equation 5.1 can be viewed as the reduced form of a selection equation

[12]The retirement literature has suggested different motives that would explain the spike of retirements at age 62, ranging from liquidity constraints that would be lifted by the Social Security pension, to work disincentives from the Social Security's earnings test, to framing effects.

[13]A similar argument applies to age 59, when defined-contribution pension holders can first access their pension.

with an unobserved, time-constant individual effect which is specified as a correlated random effect (as in Chamberlain [1980]), and where $\bar{Z}_i$ is intended to capture its correlation with observables. The vector Z_{it} includes age dummies, all other observables that enter the wage equation, plus some exclusion restrictions. $\bar{Z}_i$ includes, on top of the time averages of Z_{it}, time-invariant regressors such as education.

The selection equation is estimated by probit regression, and the results are used to generate the inverse Mills ratio, $\hat{\lambda}_{it}^R$, which is then added to equation 4.1 as an additional regressor. Wooldridge (1995) shows that a valid test of the null hypothesis of no selection is a t-statistic on the coefficient of $\hat{\lambda}_{it}^R$ in the fixed effects estimation of equation 4.1.[14]

In the multiple-index case, two selection equations of the same form as 5.1 characterize, respectively, the decisions to partially and fully retire:

$$\begin{aligned} s_{it}^{PR} &= I[Z_{it}\beta_1 + \bar{Z}_i\Pi_1 + v_{it1} > 0] \\ s_{it}^{FR} &= I[Z_{it}\beta_2 + \bar{Z}_i\Pi_2 + v_{it2} > 0], \end{aligned} \tag{5.2}$$

with v_{it1} and v_{it2} jointly normal conditional on Z_i. 5.2 is estimated by multinomial probit. Under the assumption of independent idiosyncratic error terms in the partial and full retirement equations, the two inverse Mills ratios $\hat{\lambda}_{it}^{PR}$ and $\hat{\lambda}_{it}^{FR}$ can be added to the wage equation separately to perform the selection test.

First-stage results.

Coefficients from the first-stage regressions are shown on table A.1 in the table appendix. For ease of interpretation, average marginal effects are also reported in table 3. I consider different exclusion restrictions. Column (1) shows results from a specification where an indicator that the individual has a health condition that limits his ability to work is used as an exclusion restriction. Recall that the within-individual average of the health limitation variable enters the regression as part of the vector $\bar{Z}_i$. Hence the exclusion restriction captures the effect of *deviations* from average health. This exclusion restriction is valid if these health changes do not affect wages, either because they do not influence productivity, or because employers face restrictions to reduce wages of sick workers.

The specification in column (2) considers household wealth as an alternative exclusion restriction.[15] The validity of this variable as an exclusion restriction has been questioned in the context of Heckman-type selection models of female labor supply, as it is likely correlated with workers' past

[14]No functional form assumption is required on the joint distribution of u_{it} and v_{it} in this context where we only aim to test for sample selection. The standard joint normality assumption is only required in this framework in order to correct for sample selection, once this has been detected.

[15]Health limitations are included as explanatory variables in this specification, but they are not used as exclusion restrictions -i.e. they are entered as regressors in the second stage too.

participation decisions and, hence, their attachment to the labor force and current wages. In the panel framework, however, we can control for average wealth over the sample period, to capture the correlation between lifetime wealth and the fixed individual component. The exclusion restriction is the *deviation* of current wealth from average wealth, which is arguably more exogenous.

In column (3) of table 3 I consider two exclusion restrictions that have been previously used in retirement studies to proxy for exogenous changes in labor force participation at older ages, namely indicators that the individual is over age 62 and over age 65. These are, respectively, the Social Security early and full retirement ages. Large retirement spikes at these ages have been amply documented in the literature. These are likely the most plausible exclusion restrictions, as they are part of the regulatory framework and independent of individual characteristics. The drawback is that, if they are to be excluded from the wage equation, these age dummies cannot be used to represent the age-wage profile. Therefore the second stage from this specification will make the assumption that the wage profile is well approximated by a smooth polynomial in age. Participation must also change smoothly with age, except for the institutionally-induced discrete jumps at ages 62 and 65.

The specification shown in column (4) considers the same exclusion restrictions as (3), but it allows for selection on two margins, namely partial and full retirement. The effects of observables on each form of retirement are reported separately.

As can be seen on the second panel of table A.1, all exclusion restrictions are strongly significant and of the anticipated sign. In column (2), the coefficients on both current wealth and current wealth squared are statistically significant, indicating a concave relationship between deviations from average wealth and retirement. The marginal effects for the exclusion restrictions are reported in the second panel of table 3. In column (1), having a health limitation is associated, on average, with a 14 percentage-point increase in the probability of retirement. In the second specification, a \$100k increase in wealth raises the average probability of retirement by 1.5 percentage points. Results from the third specification show that exits from full-time jobs increase by 13 percentage points for workers above age 62, and by 4 percentage points for those above 65. Comparing these results with those in column (4) we can see that crossing the Social Security retirement ages is strongly associated with transitions into full retirement, but the marginal effects are not significant in the partial retirement equation.

The main time-varying regressor is age, which, not surprisingly, turns out to be a key determinant of retirement decisions in all specifications. From columns (1) to (3) in table 3 we can see that the full-time participation profile declines steeply as workers age, while column (4) shows that there are no differential inflows into partial and full retirement with age.

The marginal effects for the variables proxying for the individual effect are shown on the third panel of table 3. Average age has a negative and strongly significant effect on retirement, indicating that later-born workers (who, given the design of the HRS panel, enter the sample at younger ages, and hence have a lower average age) are more likely to enter retirement at every age. This reflects a secular trend into early retirement, and shows that even in a narrowly-defined sample in terms of year of birth there is scope for cohort effects. Not surprisingly, the lower an individual's health type -proxied by the average number of periods he has a health condition that limits work-, the higher the likelihood that he will stop working full-time, and we can see from column (4) that both current and average health predict entrances into full retirement more strongly than those into partial retirement.

Defined contribution (DC) pension-holders are less likely to enter any form of retirement. Those with a defined benefit (DB) pension plan are also less likely to partially retire than individuals without a pension plan, but the marginal effect of the DB pension indicator is not significant in the full retirement equation. Columns (1) to (3) of table 3 show no significant effect of education on full-time participation, but we can see in column (4) that this is masking a differential effect on the two forms of retirement: college-educated individuals are 4 percentage points more likely to be partially retired, and 5 percentage points less likely to be fully retired, than individuals who have not finished high-school. For high school graduates the signs of the marginal effects are the same, but they are not statistically significant. The mother's education level, which is intended to proxy for the worker's ability, is not significant in the single-index specifications, but it is associated with a higher likelihood of both types of retirement in the multiple-index one.

Second-stage results.

Results for the second stage are presented in table 4. The coefficients on the estimated inverse Mills ratio in specifications (1) to (3) are very close to zero and not significant. The null hypothesis of no selection bias cannot be rejected in any of the three cases. In specification (4), the coefficient on the inverse Mills ratio from the full retirement selection equation is positive, while the one from the partial retirement equation is negative. However, none of the coefficients is significant. The p-values from individual t-tests on each Mills ratio are, respectively, 0.41 and 0.29. The p-value from a joint test of significance is 0.55. Once more, we cannot reject the null hypothesis of no selection on unobservables.

After the introduction of the inverse Mills ratios in the wage regressions, the age profiles remain very close to those presented in section 4. The age dummies in the first two specifications indicate a slightly increasing wage profile in the early 50's, with no significant changes thereafter (the p-value for the joint significance test of the ages dummies from years 58 to 69 is 0.21 in the first column and 0.24 in the second one). The last two columns approximate the age-wage profile with a quadratic polynomial. The age coefficients are significant, indicating a concave profile, which results from the initial increase and subsequent flattening of hourly wages. The overall wage decline for workers in their 60's, even under this restrictive functional form, is very small. Finally, there is no significant effect of current wealth and health limitations on wages. The latter result is consistent with French (2005), who also finds little evidence of wage differences by health status.

6 Is (partial) retirement a choice?

Having established that wages and earnings decline significantly only when workers transit into partial retirement, this section focuses on the nature of this transition. Are most exits from full-time jobs the result of events outside the worker's control, such as layoffs or disabling health conditions? Or do workers retire for reasons mostly unrelated to wages, such as pension incentives or a willingness to enjoy more leisure, even when they could have remained employed at their former wage rate?

The implications of the findings in this paper regarding the shape of the wage path depend crucially on the answers to these questions. To see why, suppose for simplicity that full retirement is not an option and workers are observed either in full-time or part-time jobs, with lower average wages in the latter case. If most exits from full-time jobs are involuntary, the subsequent wage drop is exogenous from the worker's perspective. The drop in hours can be rationalized in the context of the canonical labor supply model as an optimal response to the exogenous wage change. In this scenario, the usual modeling assumptions hold, and only the smoothly-declining wage profile needs to be substituted by a step-function profile, with uncertainty regarding the timing of the transition from full-time work.

If, on the other hand, most exits from full-time jobs are voluntary, then observed wages of partially-retired workers are not a good proxy for offered wages, to the extent that they could have kept their full-time job, had they chosen to. In this case, the appropriate approach to modeling labor supply must recognize that the worker faces a joint choice of wage rate and hours: he can work full time for an hourly wage w, or work part-time for a lower hourly rate.

The last scenario calls into question the conclusions of models that estimate an exogenous wage process for older workers from observed wages, and rationalize hours choices as an optimal response to a given wage rate. In particular, ignoring the joint choice of wages and hours at older ages will yield biased estimates of the intertemporal elasticity of substitution of labor supply (as shown by Chang, Kim, Kwon and Rogerson [2001]), and consequently unreliable estimates of labor supply responses to business cycles and reforms of the tax system, private pensions or Social Security. Models of income dynamics will overstate the risk facing workers if wage drops at older ages are assumed exogenous when they are not. This can affect our understanding of individuals' responses to risk and the assessment of welfare effects of social insurance programs intended to reduce this risk.[16] Retirement models will attribute too many full-time work exits to exogenously declining wages, in detriment of pension incentives or a preference for leisure (as shown by Gustman and Steinmeier [1986]).

Establishing the extent to which partial and full retirements are voluntary is not an easy task, since such pivotal decisions in a worker's lifetime are likely influenced by a myriad of motives. This

[16]See Cunha, Heckman and Navarro (2005) and Low, Meghir and Pistaferri (2010) for a discussion of the importance of the distinction between forecastable changes in wages or income from those that are the result of exogenous shocks.

is clear from the first-stage regressions in section 5, which provide an overview of the main variables that correlate with full-time work exits. Some of those, like lifetime wealth or the type of pension plan, hint at voluntary retirements that respond to incentives that are independent from wages or productivity changes. Others, such as the onset of a health condition that limits work, suggest that some retirements are triggered by factors outside the individual's control. This section reviews several pieces of evidence that, on the whole, point to a majority of retirements being voluntary decisions of workers who had the choice to remain in their full-time jobs.

Evidence from individual self-reports is presented in Szinovacz and Davey (2005), who analyze HRS questions asking respondents whether their decision to retire was voluntary. They find that 70 percent of male respondents perceive their retirement as "wanted". The remaining 30% is likely an upper bound on retirements that respond to health or work-related reasons such as wage decreases or layoffs, as some workers in the "forced" category report having to leave their jobs to care for their spouses, parents or grandchildren. In a typical retirement model, any of those reasons would be modeled as a choice resulting from the worker's comparison of foregone wages versus the cost of private care.

At the same time, some individuals may report that they have retired voluntarily whenever they had a choice to keep their full-time job, even if that would have resulted in a wage cut. Schultz, Morton and Weckerle (1998) investigate this possibility using a sample of early retirees from the HRS out which roughly 30% (70%) report having retired involuntarily (voluntarily). They find that involuntary retirees often cite poor health as the most influential factor driving their decision to retire, whereas for voluntary ones it almost always responds to their desire to "do other things" or lack of a need to work "based on adequate finances".

Apart from their subjective perceptions, HRS respondents are also asked about the specific reasons why they left their last employer. This question is asked to every person who changes jobs between waves, independently of their retirement status. I analyze the responses to this question in table 5. Out of all possible reasons for retirement, the ones shown on the first three rows (the business where the individual was working closed, he was laid off, or he was in bad health or became disabled) are most likely associated with an involuntary retirement. The remaining categories (the worker retired, had care obligations, needed to move with the spouse, etc.) are likely indicative of a voluntary retirement.

The first two columns of the table show responses of workers who have just transited from full-time work into full-retirement and partial retirement, respectively. Roughly 15% of these workers report having left the full-time job because the business closed or they were laid off. Being in poor health or disabled is a more frequent response for those who fully retired (16%) than those partially retired (4%), consistent with the results from the selection equations in the previous section. In both cases, the most frequent reason given for the recent job exit is that the worker retired (62 and 55%). Since this answer is so prevalent, table 6 probes further into the reasons behind it. All workers who report having changed jobs because they "retired" are asked whether this was following a change in employment conditions and, if so, of what type. As can be seen in the first two columns of table 6, the overwhelming majority of fully and partially retired workers report having retired following either no change in their employment conditions (80 and 75%, respectively), or a positive change, such as becoming eligible for a pension or an early retirement incentive (5 and 9%). Virtually none of these workers reports changes in their employment conditions suggesting an exogenous change in wages or in employment availability.

Finally, for comparison, the last two columns of table 5 report the reasons for job changes of workers who moved from full-time to full-time jobs, and of those who changed jobs but had already retired in the past. None of these workers counts as a transition into partial retirement in my sample. Workers switching between full-time jobs are most likely to report exogenous changes in their working conditions, such a lay-off or business closure (43%), and much less likely to indicate that they have retired (14.23%). Workers who were already partially retired and change jobs give answers that are very close to those reported in column 1, suggesting that these transitions are mostly capturing switches from a part-time job into full-retirement.

To conclude, the evidence presented in this section strongly points to a large proportion of career job exits being voluntary. The number of involuntary retirements, up to 30% of the total, is not negligible, and points to the key role of uncertainty in the working decision of older workers. However, at least 70% of full retirements and 80% of partial retirements are the result of choice. This implies that in a majority of cases the wage drop following transitions into partial retirement is endogenous.

7 Conclusions

This paper shows that the wage profile of the typical older individual is best represented by a step function: wages are flat or slightly increasing with age for as long as the worker is employed full-time. They drop discretely when he leaves the career job, and remain mostly flat thereafter. More importantly, the wage change at the transition into partial retirement is endogenous in most cases, as workers choose when to start the gradual transition into retirement.

The paper challenges three common assumptions in the empirical literature. First, that wages decline slowly with age for individuals approaching retirement, following the path of productivity. Second, that observed wages are a good proxy for offered wages in the sample of employed individuals. Third, that hours are a response to exogenous changes in wages. The results indicate that, at least at older ages, wages and hours are determined jointly, as the result of a choice between a bundle involving full-time hours and high wages, and one involving part-time hours and a lower wage rate.

The conclusions of the paper are relevant to empirical labor and macro models whose conclusions depend on the correct specification of the lifetime wage or earnings profiles. This type of models is used in very different contexts, including the analysis of labor supply changes over the business cycle, labor supply responses to tax or pension changes, the role of uncertainty in individual choices, and retirement decisions.

Finally, the conclusions of the paper are not incompatible with the prediction from the human capital model of a close association between wages and productivity. However, I find no evidence of downward wage adjustments for as long as workers are employed full-time, or of less productive workers being forced to leave their full-time jobs (with the exception of those who suffer the onset of health conditions). The findings presented here are consistent with the human capital model if productivity does not start declining before workers partially retire. On the other hand, they are also consistent with alternative theories of wage determination that do not predict a close association between productivity and wages, such as some efficiency wage models.

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Table 1: Age-Wage Profile from Different Specifications

	Dependent Variable				
	w_{it}	w_{it}	w_{it}	w_{it}^{FT}	w_{it}^{PT}
	OLS	FE	FE	FE	FE
PR=1			-0.321**		
			(0.021)		
age≥52	0.005	0.032	0.032	0.035	
	(0.035)	(0.028)	(0.027)	(0.023)	
age≥53	-0.022	0.001	0.008	0.003	
	(0.030)	(0.024)	(0.022)	(0.019)	
age≥54	0.016	-0.010	-0.005	0.011	
	(0.026)	(0.021)	(0.019)	(0.016)	
age≥55	-0.011	-0.005	0.001	0.007	
	(0.026)	(0.019)	(0.018)	(0.015)	
age≥56	0.016	0.019	0.026	0.021	
	(0.023)	(0.017)	(0.016)	(0.012)	
age≥57	-0.010	-0.002	0.006	0.010	
	(0.023)	(0.017)	(0.016)	(0.012)	
age≥58	-0.014	-0.003	0.007	-0.003	0.088
	(0.022)	(0.016)	(0.015)	(0.012)	(0.055)
age≥59	-0.027	-0.034*	-0.024	0.003	-0.114*
	(0.022)	(0.016)	(0.015)	(0.012)	(0.055)
age≥60	0.005	0.007	0.012	0.012	-0.015
	(0.022)	(0.016)	(0.015)	(0.012)	(0.050)
age≥61	-0.029	-0.006	0.012	0.001	0.091
	(0.023)	(0.016)	(0.016)	(0.013)	(0.052)
age≥62	-0.035	-0.037*	-0.005	-0.002	-0.079
	(0.026)	(0.018)	(0.017)	(0.015)	(0.046)
age≥63	-0.045	-0.033	-0.014	0.014	-0.036
	(0.031)	(0.022)	(0.020)	(0.017)	(0.048)
age≥64	-0.080*	-0.037	-0.014	-0.024	0.022
	(0.033)	(0.023)	(0.022)	(0.018)	(0.052)
age≥65	0.009	-0.006	0.020	0.014	0.050
	(0.036)	(0.024)	(0.022)	(0.023)	(0.041)
age≥66	-0.092*	-0.053	-0.036	-0.007	-0.058
	(0.040)	(0.029)	(0.028)	(0.029)	(0.039)
age≥67	0.017	-0.005	0.017	-0.044	0.052
	(0.041)	(0.031)	(0.030)	(0.035)	(0.040)
age≥68	-0.057	-0.059	-0.034	0.004	-0.078
	(0.046)	(0.038)	(0.037)	(0.041)	(0.045)
age≥69	0.006	0.041	0.036	-0.007	0.065
	(0.054)	(0.044)	(0.043)	(0.039)	(0.058)
individual-year obs.	10,028	9,537	9,537	7,165	1,800
# of individuals		2,253	2,253	1,892	583
R^2	0.208	0.06	0.15	0.01	0.02
Tests of Joint Significance (p-value):					
Age≥52-Age≥57	0.976	0.425	0.048	0.003	
Age≥58-Age≥69	0.000	0.000	0.302	0.243	0.048

NOTE. - Robust standard errors in parentheses. * indicates significance at 5%; ** indicates significance at 1%. All regressions include an intercept and a measure of the yearly unemployment rate. OLS regression includes controls for marital status, education (defined as less than high school/high school/college) and race (defined as black/hispanic/white). OLS standard errors clustered at the individual level. For fixed-effects estimates, reported R^2 is for the within regression. Dummies for ages 52 to 57 are eliminated from PT wage regression because of the low number of observations available.

Table 2: Age-Hours Profile from Different Specifications

	Dependent Variable				
	$hours_{it}$	$hours_{it}$	$hours_{it}$	$hours_{it}^{FT}$	$hours_{it}^{PT}$
	OLS	FE	FE	FE	FE
PR=1			-0.595**		
			(0.022)		
age≥52	-0.028	-0.025	-0.024	-0.016	
	(0.015)	(0.021)	(0.022)	(0.012)	
age≥53	0.021	0.039	0.051**	0.004	
	(0.013)	(0.020)	(0.020)	(0.010)	
age≥54	0.001	-0.009	0.002	0.005	
	(0.011)	(0.019)	(0.017)	(0.009)	
age≥55	-0.007	-0.016	-0.005	-0.011	
	(0.011)	(0.018)	(0.016)	(0.008)	
age≥56	-0.019	-0.011	-0.000	-0.013	
	(0.011)	(0.016)	(0.014)	(0.007)	
age≥57	-0.001	-0.005	0.010	0.001	
	(0.011)	(0.016)	(0.014)	(0.007)	
age≥58	-0.017	-0.008	0.009	-0.008	0.068
	(0.012)	(0.017)	(0.014)	(0.008)	(0.075)
age≥59	-0.001	-0.013	0.004	0.008	-0.042
	(0.014)	(0.017)	(0.015)	(0.007)	(0.081)
age≥60	-0.017	-0.020	-0.012	-0.014*	0.037
	(0.014)	(0.017)	(0.014)	(0.007)	(0.064)
age≥61	-0.008	-0.028	0.006	0.001	-0.017
	(0.015)	(0.017)	(0.015)	(0.007)	(0.066)
age≥62	-0.111**	-0.093**	-0.034	-0.020*	-0.049
	(0.022)	(0.022)	(0.018)	(0.008)	(0.061)
age≥63	-0.039	-0.072**	-0.037	-0.001	-0.109
	(0.028)	(0.027)	(0.022)	(0.009)	(0.063)
age≥64	-0.061	-0.048	-0.005	0.006	0.045
	(0.036)	(0.032)	(0.027)	(0.011)	(0.061)
age≥65	-0.021	-0.043	0.006	-0.016	0.004
	(0.039)	(0.034)	(0.029)	(0.012)	(0.057)
age≥66	-0.069	-0.053	-0.023	-0.016	0.005
	(0.044)	(0.037)	(0.032)	(0.016)	(0.054)
age≥67	0.023	0.018	0.059	0.016	0.077
	(0.045)	(0.039)	(0.034)	(0.019)	(0.056)
age≥68	-0.096*	-0.083*	-0.037	-0.036	-0.127*
	(0.047)	(0.042)	(0.038)	(0.020)	(0.062)
age≥69	-0.070	-0.075	-0.082	0.024	0.006
	(0.064)	(0.057)	(0.052)	(0.024)	(0.081)
individual-year obs.	10,028	9,537	9,537	7,165	1,800
# of individuals		2,253	2,253	1,892	583
R^2	0.12	0.17	0.37	0.02	0.02
Tests of Joint Significance (p-value):					
Age≥52-Age≥57	0.024	0.032	0.027	0.009	
Age≥58-Age≥69	0.000	0.000	0.000	0.000	0.042

NOTE. - Robust standard errors in parentheses. * indicates significance at 5%; ** indicates significance at 1%. All regressions include an intercept and a measure of the yearly unemployment rate. OLS regression includes controls for marital status, education (defined as less than high school/high school/college) and race (defined as black/hispanic/white). OLS standard errors clustered at the individual level. For fixed-effects estimates, reported R^2 is for the within regression. Dummies for ages 52 to 57 are eliminated from PT wage regression because of the low number of observations available.

Table 3: First-stage results for different specifications. Marginal Effects. (Continued on next page)

	(1)	(2)	(3)	(4)	
	Retired	Retired	Retired.	PT	Fully Ret.
Time-Varying Regressors					
age^a			0.036**	0.039**	0.040**
			(0.002)	(0.003)	(0.002)
age≥53	0.057	0.050			
	(0.038)	(0.037)			
age≥54	0.071*	0.071*			
	(0.030)	(0.031)			
age≥55	0.029	0.028			
	(0.025)	(0.025)			
age≥56	0.019	0.017			
	(0.020)	(0.020)			
age≥57	0.041*	0.040*			
	(0.018)	(0.018)			
age≥58	-0.001	-0.004			
	(0.015)	(0.015)			
age≥59	0.054**	0.057**			
	(0.016)	(0.016)			
age≥60	0.017	0.009			
	(0.015)	(0.014)			
age≥61	0.085**	0.091**			
	(0.016)	(0.016)			
age≥62	0.137**	0.132**			
	(0.017)	(0.017)			
age≥63	0.056**	0.055**			
	(0.015)	(0.015)			
age≥64	0.047**	0.047**			
	(0.015)	(0.015)			
age≥65	0.078**	0.074**			
	(0.017)	(0.017)			
age≥66	0.053**	0.053**			
	(0.018)	(0.017)			
age≥67	0.025	0.026			
	(0.018)	(0.017)			
age≥68	0.073**	0.071**			
	(0.022)	(0.021)			
age≥69	-0.003	-0.002			
	(0.023)	(0.022)			
health limit		0.138**	0.139**	-0.061**	0.194**
		(0.009)	(0.009)	(0.008)	(0.010)
current wealth^a (in 100k)			0.014**	0.011*	0.017**
			(0.004)	(0.005)	(0.004)

Table continued on next page.

Table 3: First-stage results for different specifications. Marginal Effects. (Cont.)

	(1)	(2)	(3)	(4)	
	Retirement	Retirement	Retirement	PT	Full Ret.
Exclusion Restrictions					
health limit	0.139**				
	(0.010)				
current wealth[a] (in 100k)		0.014**			
		(0.004)			
over ERA			0.128**	0.017	0.110**
			(0.011)	(0.010)	(0.012)
over FRA			0.044**	0.023*	0.021
			(0.012)	(0.011)	(0.011)
Time-Constant Regressors					
avg. age[a]	-0.014**	-0.015**	-0.017**	-0.020**	-0.017**
	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)
avg. health limit	0.130**	0.145**	0.144**	0.083*	0.189**
	(0.022)	(0.021)	(0.021)	(0.033)	(0.021)
avg. wealth[a] (in 100k)		0.005	0.005	0.005	0.005
		(0.003)	(0.003)	(0.005)	(0.004)
education = college	0.027	-0.006	-0.005	0.050**	-0.054**
	(0.015)	(0.015)	(0.015)	(0.015)	(0.015)
education = high school	0.019	0.007	0.007	0.023	-0.015
	(0.014)	(0.014)	(0.013)	(0.014)	(0.013)
mother's education	0.003	-0.002	-0.002	0.053**	0.177**
	(0.006)	(0.005)	(0.005)	(0.014)	(0.010)
pension type = DB	-0.056**	-0.067**	-0.069**	-0.097**	0.028*
	(0.014)	(0.014)	(0.014)	(0.013)	(0.014)
pension type = DC	-0.093**	-0.103**	-0.105**	-0.052**	-0.054**
	(0.016)	(0.015)	(0.015)	(0.013)	(0.014)
Obs.	16,450	16,353	16,353	16,353	16,353

NOTE. - Marginal effects reported. Standard errors obtained from 1,000 bootstrap replications clustered at the individual level to account for the panel structure of the data. * indicates significance at 5%; ** indicates significance at 1%. All regressions include an intercept and a measure of the yearly unemployment rate. Top 0.5% of observations from wealth distribution are dropped from specifications where wealth appears as regressor.

[a] Marginal effects of age, average age, wealth, and average wealth obtained from a regression including a quadratic polynomial on the corresponding variable (coefficients from this regression shown in table A.1 in the table appendix).

Table 4: Wage-Regression Equation.

	(1)	(2)	(3)	(4)
age				0.117**
				(0.030)
age^2				-0.001**
				(0.000)
age≥52	0.035	0.035	0.034	
	(0.022)	(0.022)	(0.024)	
age≥53	0.008	0.006	0.002	
	(0.021)	(0.020)	(0.022)	
age≥54	0.017	0.015	0.009	
	(0.018)	(0.018)	(0.018)	
age≥55	0.010	0.009	0.006	
	(0.015)	(0.015)	(0.017)	
age≥56	0.023	0.023	0.020	
	(0.013)	(0.013)	(0.014)	
age≥57	0.013	0.012	0.009	
	(0.012)	(0.012)	(0.014)	
age≥58	-0.003	-0.002	-0.004	
	(0.012)	(0.012)	(0.014)	
age≥59	0.008	0.006	0.003	
	(0.013)	(0.013)	(0.014)	
age≥60	0.013	0.012	0.010	
	(0.013)	(0.013)	(0.015)	
age≥61	0.007	0.006	0.000	
	(0.014)	(0.014)	(0.014)	
age≥62	0.006	0.004	-0.005	
	(0.018)	(0.017)	(0.031)	
age≥63	0.017	0.016	0.013	
	(0.018)	(0.017)	(0.018)	
age≥64	-0.021	-0.022	-0.024	
	(0.019)	(0.019)	(0.020)	
age≥65	0.017	0.015	0.012	
	(0.024)	(0.024)	(0.025)	
age≥66	-0.005	-0.003	-0.008	
	(0.030)	(0.030)	(0.030)	
age≥67	-0.044	-0.046	-0.045	
	(0.035)	(0.036)	(0.036)	
age≥68	0.007	0.006	0.003	
	(0.041)	(0.042)	(0.041)	
age≥69	-0.007	-0.009	-0.008	
	(0.040)	(0.041)	(0.040)	
health limit				-0.008
				(0.019)
wealth (in 100k)				0.003
				(0.006)
wealth (in 100k) squared				-0.000
				(0.000)
λ	0.025	0.018	-0.009	
	(0.026)	(0.024)	(0.068)	
λ^{FR}				0.032
				(0.036)
λ^{PT}				-0.070
				(0.062)
individual-year obs.	7,165 R	7,165 R	7,165 R	7,165
# of individuals	1,892	1,892	1,892	1,892

NOTE. - Standard errors (in parentheses) obtained from 1,000 bootstrap replications clustered at the individual level to account for panel structure of the data. The first and second stages are bootstrapped jointly, to obtain consistent estimates of the standard errors on inverse Mills ratios in the second stage. * indicates significance at 5%; **indicates significance at 1%;

Table 5: Reasons left last employer (in percentage points)

	i not retired in $t-1$			i retired in $t-1$
	i fully retired in t	i partially retired in t	i not retired career in t	
Business closed	5.09	4.41	18.85	5.66
Laid off/let go	9.96	12.06	23.84	11.21
Poor health/disabled	16.08	3.82	2.03	14.34
Family care	0.81	1.18	0.55	1.21
Better job	0.81	8.24	20.70	4.95
Quit	3.10	10.00	13.86	11.31
Retired[a]	61.73	54.71	14.23	38.79
Family moved	0.15	0.59	1.48	0.91
Sold (own) business	0.22	0.00	0.18	0.71
Divorce/Separation	0.00	0.00	0.00	0.10
Handed over responsibilities to other family members	0.00	0.00	0.00	0.10
Transportation; distance to work	0.00	0.29	0.74	0.30
To travel	0.00	0.00	0.00	0.20
Early retirement incentive	0.52	0.59	0.37	0.30
Financially advantageous not to work	0.15	0.00	0.00	0.30
Respondent or spouse was transferred	0.00	0.00	0.18	0.00
Other	1.40	4.12	2.95	9.60
Obs.	1,356	340	541	990

[a] See table 6 for disaggregated reasons for retirement.

NOTE: Not all individuals who are asked about the reasons why the left their last employer give an answer. The proportions in the table have been computed using nonmissing responses only.

Table 6: Reasons for retirement (in percentage points)

	i not retired in $t-1$			i retired in $t-1$
	i fully retired in t	i partially retired in t	i not retired career in t	
No change in employment situation	80.02	74.59	76.62	82.46
Wages reduced	0.12	0.00	0.00	0.29
Hours reduced	0.49	0.00	0.00	0.88
Would have been laid off	0.61	2.70	2.59	0.88
Supervisor/co-workers encouraged departure	1.36	1.62	1.29	1.17
New job duties/location	2.10	1.62	0.00	2.34
Became eligible for pension/ financial incentive	5.17	8.64	7.79	3.80
Didn't get along/unsafe working conditions	6.28	7.02	7.79	4.38
Physical condition/stress	0.74	0.54	2.59	0.71
Other	3.11	3.27	1.33	2.51
Obs.	811	185	77	342

Table Appendix

Table A.1: Coefficient from first-stage regressions. (Continued on next page)

	(1)	(2)	(3)	(4)	
	R	R	R	PR	FR
Time-Varying Regressors					
age			-0.113	-0.109	-0.095
			(0.119)	(0.179)	(0.173)
age squared			0.002*	0.002	0.003
			(0.001)	(0.002)	(0.001)
age≥53	0.305	0.281			
	(0.160)	(0.163)			
age≥54	0.275*	0.277*			
	(0.126)	(0.129)			
age≥55	0.168	0.170			
	(0.101)	(0.103)			
age≥56	0.068	0.058			
	(0.084)	(0.086)			
age≥57	0.141	0.138			
	(0.076)	(0.077)			
age≥58	0.038	0.030			
	(0.068)	(0.069)			
age≥59	0.206**	0.221**			
	(0.064)	(0.064)			
age≥60	0.101	0.075			
	(0.060)	(0.061)			
age≥61	0.322**	0.348**			
	(0.057)	(0.058)			
age≥62	0.522**	0.514**			
	(0.057)	(0.057)			
age≥63	0.224**	0.227**			
	(0.058)	(0.059)			
age≥64	0.189**	0.192**			
	(0.061)	(0.062)			
age≥65	0.308**	0.300**			
	(0.067)	(0.067)			
age≥66	0.225**	0.227**			
	(0.073)	(0.074)			
age≥67	0.095	0.103			
	(0.078)	(0.079)			
age≥68	0.314**	0.310**			
	(0.088)	(0.089)			
age≥69	-0.010	-0.003			
	(0.098)	(0.100)			
health limit		0.608**	0.612**	0.287**	1.034**
		(0.045)	(0.045)	(0.074)	(0.064)
current wealth (in 100k)			0.073**	0.063**	0.119**
			(0.016)	(0.023)	(0.024)
current wealth (in 100k) squared			-0.002	-0.001	-0.004*
			(0.001)	(0.001)	(0.001)

Table continued on next page.

Table A.1: Coefficients from first-stage regressions. (Cont.)

	(1)	(2)	(3)	(4)	
	R	R	R	PR	FR
Exclusion Restrictions					
health limit	0.604**				
	(0.045)				
current wealth (in 100k)		0.074**			
		(0.016)			
current wealth (in 100k) squared		-0.002*			
		(0.001)			
over ERA			0.490**	0.496**	0.733**
			(0.046)	(0.072)	(0.068)
over FRA			0.169**	0.223*	0.222*
			(0.062)	(0.093)	(0.087)
Time-Constant Regressors					
avg. age	-0.067**	-0.073**	0.494**	1.245**	0.295
	(0.006)	(0.006)	(0.173)	(0.283)	(0.265)
avg. age squared			-0.005**	-0.011**	-0.003
			(0.001)	(0.002)	(0.002)
avg. health limit	0.558**	0.645**	0.642**	0.390**	1.104**
	(0.067)	(0.068)	(0.068)	(0.112)	(0.098)
avg. wealth (in 100k)		0.030*	0.031*	0.042*	0.036
		(0.012)	(0.012)	(0.018)	(0.019)
avg. wealth (in 100k) squared		-0.000	-0.000	-0.000	-0.000
		(0.000)	(0.000)	(0.000)	(0.000)
education = college	0.110**	-0.052	-0.046	0.153*	-0.197**
	(0.035)	(0.037)	(0.037)	(0.059)	(0.055)
education = high school	0.074*	0.012	0.015	0.097	-0.024
	(0.034)	(0.034)	(0.034)	(0.056)	(0.050)
mother's education	0.011	-0.012	-0.013	0.042*	-0.054**
	(0.013)	(0.013)	(0.013)	(0.021)	(0.019)
pension type = DB	-0.183**	-0.238**	-0.244**	-0.565**	-0.155**
	(0.034)	(0.035)	(0.035)	(0.054)	(0.050)
pension type = DC	-0.378**	-0.436**	-0.443**	-0.609**	-0.565**
	(0.037)	(0.037)	(0.037)	(0.057)	(0.054)
Obs.	16,136	16,040	16,040	16,040	16,040

NOTE. - Marginal effects reported. Standard errors obtained from 1,000 bootstrap replications for probit regressions and 250 replications for multinomial probit regression. * indicates significance at 5%; ** indicates significance at 1%. All regressions include an intercept and a measure of the yearly unemployment rate. Top 0.5% of observations from wealth distribution are dropped from specifications where wealth appears as regressor.