Revisiting the Hump-Shaped Wage Profile

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January 2018

Abstract

This paper shows that the wage path of the typical individual does not decline smoothly in the years before retirement, as implied by the popular hump-shaped specification. Instead, wages are flat for full-time workers, and wage drops are only observed for those who transit into part-time work before fully withdrawing from the labor force. The paper tests the implications of three alternative models of retirement that can generate the observed wage profile. The results of these tests are used to characterize the offered wage profile at older ages and the forces driving retirement transitions.

JEL Codes: J31, J26

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[†]An earlier version of this paper was circulated under the title "Wage and Earnings Profiles at Older Ages: Implications for the Estimation of Labor Supply Elasticities". I have benefited from insightful comments from Orazio Attanasio, Richard Blundell, Eric French, Michael Hurd, Hamish Low, Robin Lumsdaine, Maurizio Mazzocco, Lars Nesheim, Franco Peracchi, Andrew Samwick, and participants at several seminars and conferences. All errors are my own.

1 Introduction

The age profile of wages, also known as the *deterministic* or *predictable* component of wages, is a key input in empirical labor and macroeconomic models. Measures of income uncertainty, estimates of the intertemporal elasticity of substitution of labor supply, and assessments of the reasons for retirement depend on the correct specification of the age-wage path. But how does this profile look like? The most common specification found in the literature assumes that individuals face an inverted-U-shaped wage path that peaks in middle age and declines smoothly thereafter. The justification for this assumption goes back to the human capital model of Becker (1993), Ben-Porath (1967), and Mincer (1974), which predicts a flattening off and eventual decline of productivity as workers approach retirement age. To the extent that wages reflect productivity, this translates into an hump-shaped wage profile.

This paper first shows that the empirical evidence does not support the downward-sloping wage profile at older ages implied by the hump-shaped specification. Instead, the real hourly wage of the typical male over age 50 increases slightly with age for as long as he is employed full time. Two-thirds of workers transit directly from full-time work into full retirement. For these workers, no wage declines are observed at any age. The remaining one-third of workers transit into retirement more gradually, and go through at least one period of part-time work before withdrawing from the labor force -in what follows, I refer to this transitional period as *partial retirement*. The transition from full-time to part-time work is associated with a one-off 34% hourly wage drop. Thereafter, wages once more remain flat. Thus, the wage path followed by workers who partially retire is best represented by a step function. Similar results are obtained for earnings, with the drop upon transition into part-time work equal to 73% in this case. The downward-sloping wage and earnings profiles often estimated in the literature are a result of averaging over an increasing proportion of part-time employees as workers age. These average profiles are not a good proxy for the wage and earnings paths followed by the typical individual.

Having established that wage drops at the end of the working career are driven by transitions into part-time work, the next step is to study the nature of these transitions. This paper proposes and tests 3 alternative models of retirement that differ in the forces driving retirement decisions and the underlying process for offered wages. In the *self-selection* model, offered wages follow a smooth, downward trend at older ages and transitions into partial and full retirement are a response to declining wages. Although wages fall on average, different workers are hit by different idiosyncratic shocks, and those who receive positive wage shocks are more likely to remain in full-time employment. Because of this positive self-selection bias, the observed wage profile for full-time workers understates the rate of decline of offered wages at older ages.

In contrast with the self-selection model, neither of the other two models predicts a wedge between observed and offered wages. These next two models differ on the determinants of retirement transitions. In the *involuntary retirement* model, factors outside the worker's control trigger separations from the full-time job with increasing probability every year. In the *voluntary retirement* model, full and partial retirement transitions arise as optimal choices for workers who compare the relative value of work versus leisure. At every age, workers choose between a full-time job with a full-time wage w, a part-time job with wage θw, $\theta < 1$, and full retirement. Transitions into partial and full retirement are driven by factors which, unlike wages, change as workers age. These may operate through preferences (e.g., increasing taste for leisure) or the budget constraint (e.g., work disincentives after certain ages from the social security rules). In this model, wages and hours are chosen simultaneously, in sharp contrast to the canonical labor supply model, which posits that participation and hours choices are made in response to an exogenous wage rate.

The empirical evidence rejects the model of *self-selection*, indicating that there is no difference between offered and observed wages. Next, I provide evidence that between 20 and 30% of transitions out of full-time work are driven by events out of the worker's control such as health events or firm closures. The substantial number of involuntary transitions underscores the key role of uncertainty in labor supply decisions of older workers. More importantly, the remaining 70 to 80% of full and partial retirement transitions are voluntary. This result indicates that the vast majority of retirements arise as the optimal choice for workers who have the option to remain in full-time employment with their full wage, but choose instead to trade more leisure for a lower hourly wage. To summarize, the results show that the correct specification for the offered wage profile is flat in age, and that wage drops at the time of transition into partial retirement are endogenously determined for the majority of workers.

The implications of these findings are wide ranging because key results in the labor supply literature depend on the correct specification of the age-wage profile. For example, Casanova (2013) shows that estimates of the intertemporal elasticity of substitution of labor supply are upward biased when the the smoothly declining average wage profile is used as a proxy for the flat offered wage profile at older ages. In the literature estimating income uncertainty it is also common to approximate the predictable component of earnings or wages with a hump-shaped specification (see Hall and Mishkin (1982), Hubbard, Skinner and Zeldes (1994) or Floden and Linde (2001)), and to implicitly assume that deviations from this path are a result of shocks. This approach may lead to overestimation of income risk 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 age-wage profile is also a pivotal component of retirement models. If hours and wages are determined simultaneously, modeling retirement as the optimal response to an exogenously declining wage profile will understate the role of a worker's taste for leisure or financial incentives in triggering retirement transitions. Finally, different theories of wage determination have different implications concerning the relationship between productivity and wages as workers age. The hump-shaped wage profile is often justified on the grounds that it is consistent with the human capital model. The wage profile described 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 flat wage profile for full time workers is consistent with alternative wage theories such as Lazear's (1979) model of deferred compensation.

This paper builds on the literature on age-wage and age-earnings profiles, dating back to Hurd (1971), Lillard and Willis (1978), Mincer (1974), and Welch (1979). Early cross-sectional studies find wage and earnings declines setting in as early as age 40 (see Willis (1985) for a survey). Longitudinal studies such as Honig and Hanoch (1985b), Johnson and Neumark (1996), and Rupert and Zanella (2012) find that wages peak later, and observe the sharpest declines around the social security ages.[1] A few studies have pointed out that transitions into part-time work at older ages are often accompanied by significant hourly wage cuts (Aaronson and French (2004), Hurd (1996), and Gustman an Steinmeier (1985)). Finally, Murphy and Welch (1990, 1992) estimate earnings profiles for the sub-sample of full-time workers. Consistent with the findings in this paper, they observe no or very small wage declines as workers age.

The paper proceeds as follows. Section 2 describes the data and the cross-sectional

[1] Rupert and Zanella obtain mixed results regarding wage declines for older workers, which are present for some but not all of the cohorts they study. For the cohort that is the focus of their empirical analysis, they find declines in the fixed effects wage profile estimated without controls for retirement status –consistent with Honig and Hanoch, Johnson and Neumark, and this paper– for the college-educated group. They do not find wage declines for the group without a college education. It is worth pointing out that the PSID has the advantage of following individuals throughout the whole life cycle, but may not be the most appropriate data source for the study of wage profiles of older workers, given that sample sizes at these ages are extremely small.

profiles of the variables of interest; section 3 presents estimates of wage, hours, and earnings paths for the average individual; section 4 characterizes the offered wage profile and tests alternative models of retirement; and section 5 concludes.

2 Data and Variable Definition

A. Sample Description

The data used in the analysis are drawn from the University of Michigan Health and Retirement Study (HRS), a panel dataset of adults over age 50 and their spouses. Other longitudinal datasets, such as the PSID, collect information on older workers, but sample sizes are small for this age group. The HRS provides a large sample size and information on an extensive set of covariates, including employment status, job history, income, pensions, net worth, and demographic characteristics. HRS respondents are interviewed for the first time in 1992, and subsequently every two years. I use the 9 survey waves available up to 2008.[2]

I apply several restrictions to the data to produce a sample of workers that are relatively homogenous and have not yet started the retirement process. First, I limit the sample to men born between the years 1931 and 1941, which constitute the core HRS cohort; 4,895 HRS respondents fit this criterion. Second, only workers who are employed full time by the time they enter the panel are kept in the sample. This excludes 390 respondents who are partially retired in the first panel wave, and 1,152 who are fully retired. The reason for this restriction is that the wage change upon separation from the full-time job, which is the focus of the paper, is not observed for these individuals.[3] Third, because of concerns regarding the measurement of hourly wage rates for the self-employed, 1,015 individuals who report being self-employed for at least one period are also dropped. Appendix A.2 discusses the sensitivity of the results to this restriction. Finally, observations for a particular year are excluded if the hourly wage is lower that \$3 or higher than \$70 (in 2011\$), the number of hours worked cannot be established, or the individual is younger than 51 or older than 67, due to small sample sizes outside those ages.[4] The final sample consists of 2,318 individuals

[2] Where possible, HRS variables are extracted from the RAND HRS data file. For a description of this dataset, see St.Clair et al. (2010). A supplementary issue of the *Journal of Human Resources* (1995, vol. 30) provides more details about the HRS.

[3] See appendix A.1 for a description of the observable characteristics of individuals who are retired by the time they enter the HRS panel.

[4] The bounds on hourly wages are intended to minimize the effect of outliers. Rows 1 to 3 of table 6 show

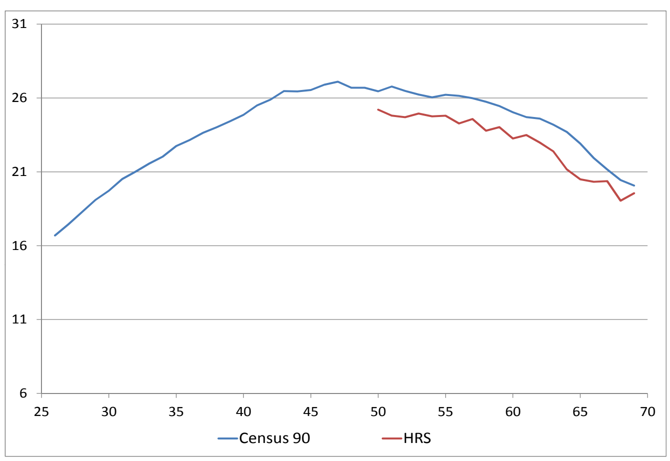
and 12,036 person-year observations.

B. Cross-Sectional Wage and Hour Profiles

This section describes the cross-sectional profiles of the variables of interest. Due to the age structure of the HRS sample, these profiles span the last years of the working life only. For comparison, lifetime average profiles obtained from the census are also shown.[5]

The focus of the paper is on the real hourly wage (expressed in 2011$ using the CPI). The first series in figure 1 shows the lifetime hourly wage profile, constructed using census data. This profile is hump shaped. It increases in the early stages of the working career, flattens out for workers in their 40s, and eventually declines from about age 50 until retirement. The inverted-U shape of the wage profile means that the well-known hump in lifetime earnings – described by Lillard and Willis (1978), Mincer (1974), Welch (1979), and Heckman, Lochner and Todd (2006), among many others– is not purely a reflection of declines in hours worked during the pre-retirement years. The second series of figure 1 shows the hourly wage profile for the pooled HRS sample. Due to differences in the definition of hourly wages –the census hourly wage rate includes commissions, cash bonuses, tips, and income from secondary jobs, while the HRS variable does not- the two series do not overlap. Nevertheless, the HRS series tracks the rate of decline in the census remarkably well.

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



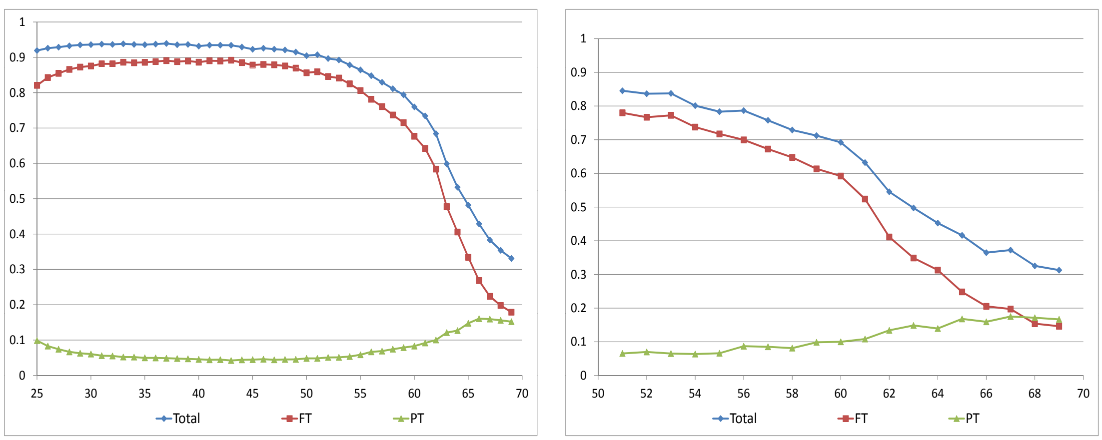
NOTE. - Hourly wages deflated to 2011 dollars using the CPI. Hourly wages lower than $3 and higher than $70 in real terms excluded. Sample includes all workers.

that using an upper bound of $50, $100, or removing it altogether hardly affects the main results. This is not surprising, as only 2% of observations in the sample have wages above $70, and just below 6% have wages above $50.

[5] Data from the 1990 US census are used. For a description of the census sample, see Ruggles et al. (2010).

The other key variable in the empirical analysis is participation status, which is used to determine whether a worker is partially retired. Work status is defined as a function of the number of hours worked per week at the current job. Workers are classified as working full time (part time) if they work at least (less than) 35 hours per week.[6] The left panel of figure 2 shows life-cycle participation profiles from the census. Two main trends are noteworthy. First, participation rates begin to decline for workers in their late 40s, at approximately the same time as hourly wages. Second, part-time participation increases threefold from age 50 to age 65. Within the sample of individuals that are employed, the rise in part-time work is even more remarkable, with the proportion of part-time workers rising from 4% to 28% between those ages. The right panel of figure 2 shows participation profiles for the pooled HRS sample. The same key trends -a rapid decline in total and full-time participation after age 50, and a simultaneous increase in part-time work- are apparent for these data.

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



C. Definition of Partial Retirement

A worker is classified as partially retired the first time he transits from full-time into part-time work. Partial retirement is modeled as an absorbing state, that is, workers who go back to full-time work are still considered to be partially retired. This modeling assumption interprets the first exit from full-time work as a signal that the worker has entered the transitional period leading to full retirement. In practice, only a small proportion of partially retired workers return to full-time work. As a result, the definitions of partial retirement and

[6]Using different definitions of full-time work, part-time work and inactivity, such as classifying as full-time workers those who work more than 30 hours per week, and considering those who work less than 2 or 5 hours per week as inactive, does not affect the results.

part-time work overlap closely, and in the rest of the paper I will refer to both interchangeably. According to this definition, 29% of workers in the sample are classified as partially retired in at least one period.

The measure of partial retirement used here is relatively parsimonious, requiring information on full- versus part-time status only. It will be easy to replicate in other contexts, given that most datasets that measure wages also provide information on the number of hours worked.[7]

3 Wage and Earnings Profiles

In this section, I first characterize the hourly age-wage profile for the typical male worker in the last stages of his working career. Next, I characterize the hours profile for the same worker. Combining the two sets of results, I construct the earnings profile for the average male over age 50.

A. Age-Wage Profile

I assume the following specification for the conditional mean of log wages:

$$w_{it} = W(Age_{it}) + X_{it}\beta + u_{it}, \tag{1}$$

where w_{it} is the logarithm of the real hourly wage of individual i in period t, W some function of age, X_{it} a vector of other observable characteristics, and u_{it} the error term.

Equation (1) can be interpreted as a standard Mincerian regression where wages, rather than earnings, are used as the dependent variable. The focus on wages allows to separate earning potential from hours choices, which are discussed separately below. 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 is captured by the fixed effect, while experience is embedded in the age effect.[8]

[7]Other studies have proposed somewhat different definitions of partial retirement. As part of the robustness checks, I re-run the analysis using 3 alternative definitions. The results are discussed in appendix A.2

[8]Using age, rather than potential experience, can confound the interpretation of the results if experience-wage profiles are not parallel across schooling levels over the age range covered in the sample. To investigate this possibility, I have run the main wage regressions separately by education category, that is, keeping potential experience constant. The results, shown in columns 7 to 9 of table 6, show that wage profiles are flat, and therefore parallel, for workers over age 50 in the 3 education groups considered. It is interesting to compare these results with those in Heckman, Lochner and Todd (2006), who reject parallel *average* earnings

I begin by estimating equation (1) by OLS, pooling all available wage observations. The results are shown in the second column of table 1. The implied wage profile, displayed as the solid grey line in figure 3, trends downwards for workers in their mid 50s and older. Specifically, the estimated coefficients imply a 3% decline in wages between ages 51 and 60 –which is not statistically significant according to the joint significance tests shown at the bottom of table 1–, followed by a strongly significant 22% decline between ages 60 and 67. The rapid wage decline at older ages is consistent with the hump-shaped lifetime wage profile described in the previous section.

In the literature, the wage profile of interest is often the individual wage profile, that is, the one that measures returns to age for the average worker. The OLS profile estimated above measures something different, namely the mean wage for the sample of individuals that remain employed at every age. It is clear that, in general, these two profiles need not overlap. And yet it is common practice to use the OLS wage profile as a proxy for the individual one, particularly when working with cross-sectional datasets, such as the CEX, that preclude following workers over time.[9] This may be a reasonable approach for individuals in their prime working years, as the vast majority are employed full time. During the retirement period, however, the characteristics of the average worker need not be stationary. The possibility that workers with a high unobserved wage component (e.g., those with high IQ or unobservable skills) retire earlier or later than the rest is of particular concern, as this would generate a correlation between the fixed component of the error term and age in equation (1), biasing the OLS age profile.

To address this concern, I next estimate equation (1) by fixed-effects.[10] Coefficients from this regression are presented in the second column of table 1. The results are also shown graphically as the dashed line in figure 3. The fixed-effects profile is mostly flat for workers in

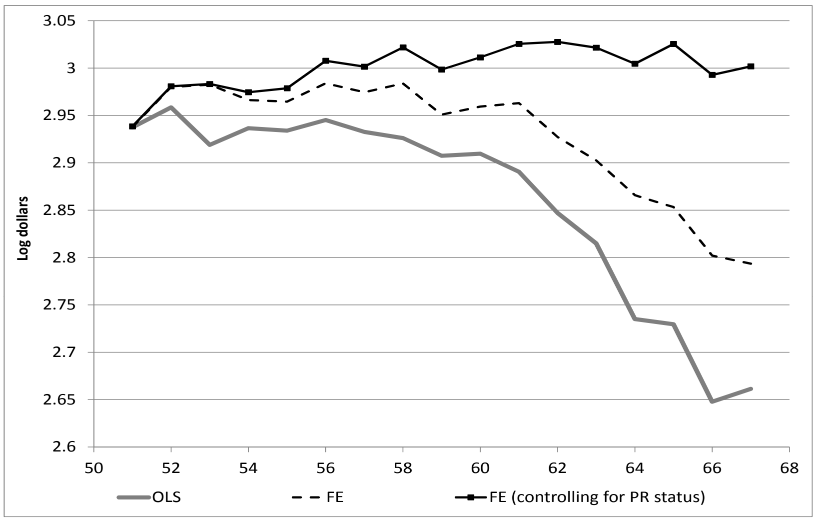
profiles. Several mechanisms can generate diverging education-specific average earnings profiles when the underlying *individual* wage profiles are flat. The most salient ones are differences in retirement propensities across education categories, and a full/part-time wage gap that is education-specific. Both mechanisms are present in the data. Table 3 shows that (partial) retirement probabilities vary with educational attainment. Table 6 shows that the wage gap between full-time workers and partial retirees increases monotonically with education.

[9]See, for example, Gomes, Kotlifoff and Viceira (2008), Guvenen (2009), Heathcote, Storesletten and Violante (2010), Kopecky (2011), and Low (2005) for OLS profiles of hourly wages or hourly earnings. Ghez and Becker (1975), Gourinchas and Parker (2002), Heckman (1976), Heckman, Lochner and Todd (2006), Krueger and Kubler (2006), Scholz, Seshadri and Khitatrakun (2006), and Storesletten, Telmer, and Yaron (2004) estimate OLS earnings profiles. As shown below, the conclusions regarding individual wage profiles extend to individual earnings profiles.

[10]It is worth noting that the fixed-effects estimator is not consistent if individuals retire as a result of wage shocks. The issue of self-selection is addressed in the next section.

their 50s, and declines rapidly from age 60 onwards. Specifically, the estimated coefficients imply a borderline significant 2% increase in real hourly wages between ages 51 and 60, followed by a strongly significant 17% decline between ages 60 and 67. The rate of decline is attenuated with respect to the OLS profile, but hourly wages still fall noticeably over the retirement period.[11]

Figure 3: Log Wage Profiles for Different Specifications



NOTE. - Series constructed using coefficients from columns 1 to 3 in table 1.

The fixed-effects profile has been used as a proxy for the individual one in studies where panel data are available.[12] Although an improvement with respect to OLS, it is still an *average* profile, measuring mean wage growth for employed individuals at every age. The increasing prevalence of part-time work during the retirement period naturally raises the question of whether average wage declines mask different wage growth rates between workers who remain in full-time employment and those who switch to a part-time job. Put differently, does the hourly wage of the typical worker decline smoothly over the retirement years, as suggested by the OLS and fixed-effect profiles? Or are declines in average wages driven by the increasing proportion of workers transiting into less well paid, part-time jobs during the retirement period?[13]

[11] Honig and Hanoch (1985b), Johnson and Neumark (1996) and Rupert and Zanella (2012) also find that the OLS profile declines more rapidly than the fixed-effects one.

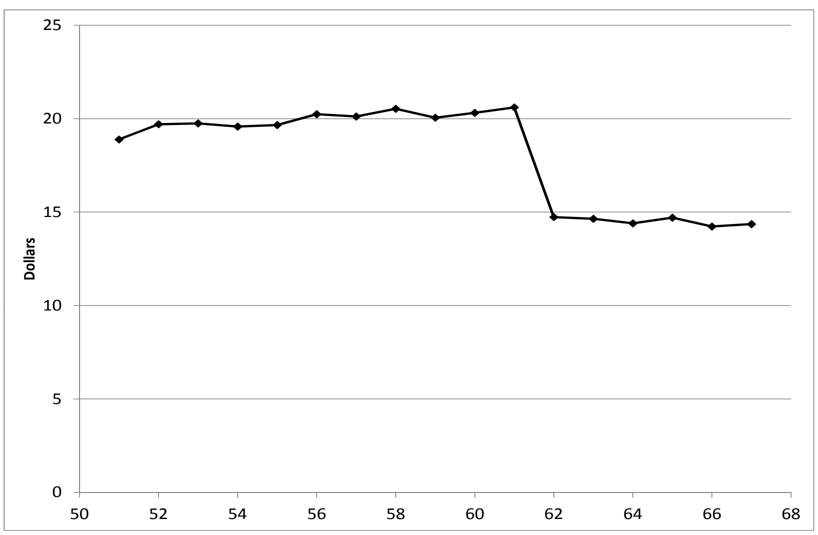
[12] See, for example, Baker (1997).

[13] The well-known fact that part-time workers earn lower hourly wages is confirmed in figure 5 below. Lower wages for partially retired workers can arise as a result of loss of tenure upon separation from the career job, a part-time wage penalty, or differences in job characteristics if partial retirees transit into occupations that

To answer this question, I re-estimate equation (1) adding as an additional regressor an indicator that is equal to 1 if the individual is partially retired. Since partial retirement is modeled as an absorbing state, its coefficient is identified from individuals who transit from full-time into part-time work for the first time. Estimates from this regression, reported in the third column of table 1, show that the transition into partial retirement is associated with a 34% decline in hourly wages. The estimated age-wage profile is shown as the solid line in figure 3. The contrast with the OLS and fixed-effects wage series is striking. After controlling for retirement status, the wage profile no longer declines with age. Wages now increase by 7% between ages 51 and 60, and by 0.7% from 60 to 67. The p-value for the test of joint significance of the age dummies spanning the retirement years is 0.70, indicating that a flat wage profile cannot be rejected for the late-career period.

The last set of results implies that average wage declines at older ages are fully driven by transitions into partial retirement. The wage path for the typical individual over age 50 is flat or increases slightly with age as long as he remains employed full time. For those who transit directly into full retirement, no wage declines are observed. Workers who partially retire, on the other hand, experience a discontinuous wage drop upon separation from the full-time job.

Figure 4: Predicted Hourly Wage Profile



NOTE. - Series constructed using coefficients from column 3, table 1, assuming that the transition into partial retirement takes place at age 62.

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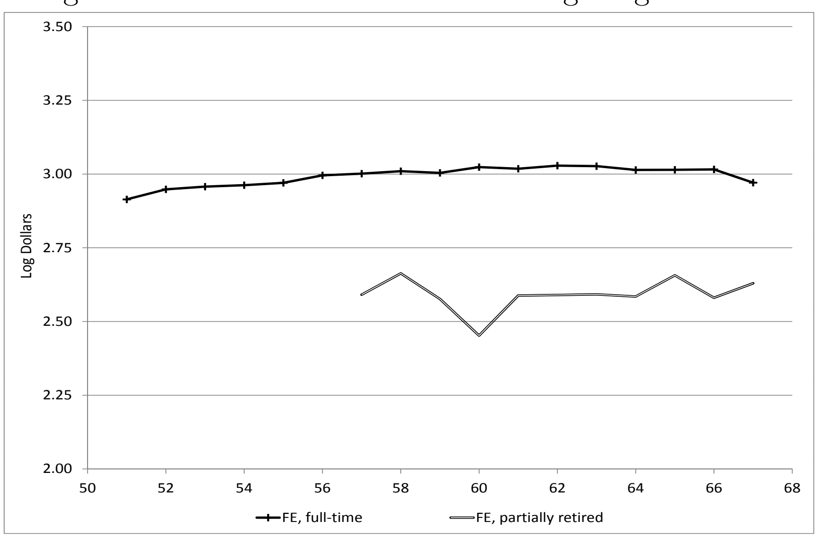
require less effort or responsibility. Distinguishing among the different mechanisms is beyond the scope of this paper. See Johnson and Neumark (1996) and Ruhm (1990) for further discussions of the wage and job characteristics of partial retirees.

Figure 4 illustrates, as an example, the predicted wage profile for an individual who transits into partial retirement at age 62. The resulting step-function profile is very different from the smoothly declining average wage profile discussed earlier. It is clear that the average wage profile is not a good proxy for the wage path followed by the typical individual.

To further illustrate how the smooth, downward-trending wage profile emerges, it is useful to consider the results of separate regressions for the subsample of non-retired workers –the *full-time* profile, shown on the left series in figure 5– and of partially retired workers –the *part-time* profile, shown on the right series in figure 5. Coefficients for these regressions are presented on the last two columns of table 1. Joint significance tests show no evidence of wage declines for full-time workers. For partially retired workers there are both positive and negative wage changes at older ages, but no discernible trend. When not controlling for partial retirement status, the resulting age-wage profile is a weighted average of these two series. The increasing weight on part-time wages as workers age and more of them become partially retired results in a downward slope.

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



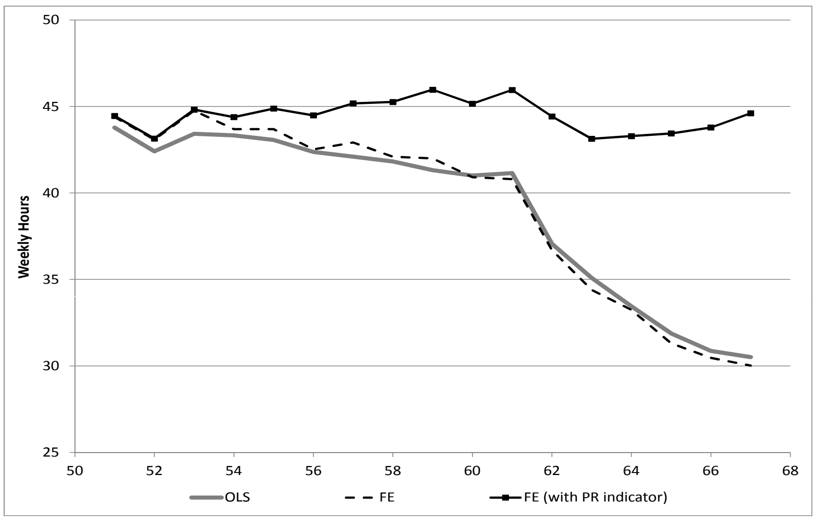
NOTE. - Series constructed using coefficients from columns 4 and 5 in table 1.

B. Age-Earnings Profile

Because structural models of consumption and saving often use the earnings profile –rather than the hourly wage profile– as an input, this section characterizes the earnings path of the typical worker over age 50. To do so, I proceed in two steps: I first analyze the (weekly) hours profile, and then combine the results with those already presented for hourly wages to characterize the (weekly) earnings profile, constructed as the product of the two.[14]

I use the same specification for the hours profile as I did for hourly wages. I estimate equations analogous to (1) using log weekly hours as the dependent variable. As before, I compare results from OLS and fixed-effects regressions estimated by pooling observations for full-time and partially retired workers to those obtained from a fixed-effects regression that controls for partial retirement status. The coefficients are presented in columns 1 to 3 of table 2. The implied hour profiles (in levels) are plotted in figure 6.

Figure 6: Hours Profiles for Different Specifications



NOTE. - Series constructed using coefficients from columns 1 to 3 in table 2.

The OLS and fixed-effect profiles are nearly identical in this case, suggesting that unobserved individual characteristics associated with a higher propensity for early retirement are uncorrelated with the number of hours worked. Once the the partial retirement indicator is added to the regression, it becomes clear that the smoothly-declining average hours profile masks a sharply discontinuous labor supply path at the individual level. This is supported by two main findings. First, the transition into partial retirement is associated with a 64%

[14] I do not use the yearly earnings measure from the HRS in this analysis because the earnings drop associated to partial retirement cannot be inferred from this measure. The HRS collects earnings information for the whole year before the time of the interview. For workers who partially retired during that year, the earnings measure includes some earnings obtained while employed full time. Because information on hourly wages and hours during the pre-retirement period is generally not available, it is not possible to accurately divide yearly earnings between pre- and post-retirement jobs. As a result, estimates of the earnings profile using yearly earnings as the dependent variable will understate the drop in earnings associated to the partial retirement transition.

drop in hours worked. Second, most of the decline in average hours at older ages is driven by transitions into partial retirement. Average hours, represented by the dashed line in figure 6, decline steeply with age, dropping by close to 8% from ages 51 to 61, and by a further 31% from ages 61 to 67. By contrast, the hour profile obtained after controlling for transitions into partial retirement, represented by the solid line in figure 6, shows no decline in hours before age 60, and a small 1% decline from 61 to 67.[15] These two findings are inconsistent with a gradual reduction of hours worked in the years leading to full retirement. Instead, labor supply adjustments are concentrated around retirement transitions. Hours drop from full time to 0 for workers who transit directly into partial retirement. For workers who partially retire, there is a further discontinuous drop in hours, illustrated on the left panel of figure 7 for a hypothetical worker who becomes partially retired at age 62. As was the case for wages, the average hour profile is not a good proxy for the labor supply path of the average individual.

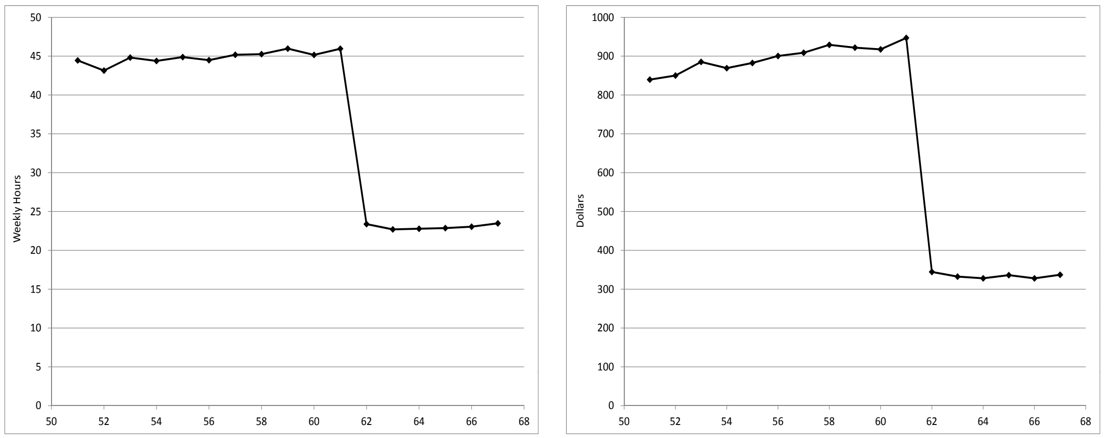
One final point worth noting is the acceleration in the downward trend of the average hour profile around the social security early and full retirement ages (62 and 65, respectively). The coefficients in column 2 of table 2 show an 11% decline in average hours at age 62, and a further 6% decline at age 65. Once the transition into partial retirement is accounted for, the coefficient on the full retirement age dummy falls to zero and becomes insignificant. The coefficient on the early retirement dummy remains significant, but its magnitude is reduced by more than two-thirds. These results confirm previous findings indicating that a large number of transitions into partial retirement take place at the social security ages, likely as a response to incentives from the social security rules (See Aaronson and French (2004), Blau (2008), French (2005) and Rust and Phelan (1997)). Since these rules are exogenous from the point of view of the individual, in the next section I will use them as exclusion restrictions for partial and full retirement transitions.

Finally, the results for wages and hours can be combined to offer some insight on the shape of the earnings profile. The right panel of figure 7 shows predicted weekly earnings for an individual who transits into partial retirement at age 62. 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.

The definition of labor earnings used in this analysis does not include bonuses, tips, or

[15]The age dummies spanning the social security ages are jointly statistically significant, suggesting that some hour adjustments precede the retirement decision. This is confirmed by the regression run using only the sample of full-time workers (column 4 of table 2), which shows a 2% hour drop at age 62.

Figure 7: Predicted Weekly Hour (left) and Weekly Earnings (right) Profiles.



NOTE. - Left series constructed using coefficients from column 3, table 2. Right series constructed combining coefficients from column 3 in tables 1 and 2. Both series assume that the transition into partial retirement takes place at age 62.

income from secondary jobs. It is conceivable that these may be adjusted before partial retirement. However, in view of the evidence in figure 7, it seems safe to conclude that the drop in earnings upon transition into partial retirement accounts for the vast majority of earnings declines observed at older ages.

4 Offered Wages

The previous section characterized the *ex post* wage path followed by a typical individual. This path is conditional on labor supply decisions: it shows that partial retirement is accompanied by large wage drops, but it is not informative about the hourly wage partial retirees would receive *were they to remain employed full time*. This section characterizes the offered wage profile, which is not confounded by hour choices.

The step function wage profile described in the previous section is consistent with at least three models of retirement. These differ in the forces driving the retirement decision and the underlying process for offered wages. In order to characterize the offered wage profile, I test the empirical implications of the three models to determine which of them is compatible with the data.

In the first model, which I refer to as the *self-selection model*, offered wages trend downwards at older ages; that is, the lifetime offered wage profile is hump-shaped. Becker (1993), Ben-Porath (1967) and Mincer (1974) provide the microfoundations for this model by assuming that offered wages are a function of an unobserved variable, such as productivity, that declines smoothly with age. To the extent that individuals allocate more work hours to periods of higher wages, they will switch from full-time work into partial or full retirement once their wage falls below some threshold. Thus, in this framework, transitions into partial and full retirement at older ages are a response to declining offered wages. The key testable implication from this model is a discrepancy between offered and observed wages. Recall that the evidence presented in the previous section shows that observed full-time wages are flat or increase slightly with age. For this to be compatible with declining offered wages, individuals who receive positive wage shocks must be more likely to remain in full-time employment. Hence, this model implies a positive self-selection bias.

Alternatively, if there is no selection, offered wages are the same as observed wages, and the offered full-time wage profile is nondecreasing in age –as shown in figure 5. In this case, we can broadly define two retirement models depending on whether retirement transitions are voluntary.

In the *involuntary retirement model*, transitions out of full-time work are due to factors outside the worker's control. For simplicity, suppose that they are driven by a discrete Markov process that the individual takes as exogenous. In each period workers are hit by a shock leading to job separation with positive probability (e.g., they become disabled or are fired). In this framework, while the *ex-post* wage profile for full-time workers is flat, the *expected* wage profile for an individual approaching the end of his career declines smoothly as long as the probability of involuntary job separation increases continuously with age.

Finally, in the *voluntary retirement model* offered wages are flat and transitions into partial and full retirement are voluntary, i.e., driven by the relative preference for leisure rather than an exogenous shock that prevents work. In a model of voluntary transitions, workers evaluate different bundles of wages and hours (e.g., wage $w_t > 0$ for full-time work, wage θw_t, with $0 < \theta < 1$, for part-time work, and no wage for full retirement), and effectively choose the two variables simultaneously. In the absence of changes in offered wages with age, other variables must be responsible for shifting the relative value of leisure and generating transitions into part-time work and out of the labor force. These variables can operate through preferences (e.g., increasing taste for leisure) or through the budget constraint (e.g., work disincentives after certain ages from the social security rules). In this model, retirement transitions are not a response to declining wages. Rather, wage and hour drops simultaneously determined when workers choose to transit from full-time into part-time work.

Each of these 3 models of retirement has empirical implications that can be tested. The remainder of the section examines which of them are compatible with the data.

A. Self-selection into Retirement

Recall that the results presented in section 3 were obtained estimating equations of the following type:

$$w_{it} = f_i + W(Age_{it}) + X_{it}\beta + u_{it} \tag{2}$$

There is a concern that estimates of (2) obtained using the sample of workers will understate the decline in wages at older ages if individuals who receive negative shocks are more likely to transit into retirement. The sufficient condition for consistency of the estimates presented in section 3 can be stated as follows:

$$E(u_{it}|f_i, Age_i, X_i, s_i) = 0, \tag{3}$$

where s_{it} is a selection indicator equal to 1 if individual i's wage is observed in period t, and $s_i = (s_{i1}, ..., s_{iT})'$. To test whether this condition is violated, I follow the procedure described in Wooldridge (1995), which extends Heckman's (1979) two-step estimator to a panel-data context.

The equation of interest is the age-profile of offered wages for full-time work, that is, the wage a worker would receive were he to choose to work full time. This is given by:

$$w_{it}^{FT^*} = g_i + A(Age_{it}) + X_{it}\delta + \varepsilon_{it}, \tag{4}$$

where w^{FT^*} is the logarithm of the real offered full time hourly wage, A is some function of age, X a vector of observables, g an individual-specific effect, and ε an idiosyncratic error term.

Two alternative specifications for the selection process are considered. The first one pools partial and full retirement together: The selection indicator d^R is equal to 1 if the individual is not working full time.[16] It is determined by the following probit equation:

[16] Notice that the selection indicator is defined in the opposite way to that used for s in equation (3) above. s indicates selection *into* the sample of workers, and d^R selection *out* of it. This slightly counter intuitive choice is made for comparability with results from the second specification for the selection process, discussed below.

$$d^R_{it} = I[Z_{it}\beta + \bar{Z}_i\Pi > -v_{it}],$$
$$v_{it}|Z_i \sim \text{Normal}(0,1) \tag{5}$$

where $Z \supset X$ is a vector of observables that contains all the variables in the vector X plus the variables used as exclusion restrictions, and $\bar{Z}_i$ a vector of within-individual averages.[17] v_{it} is a normally distributed error term.[18] Full-time wages are observed for individuals who are neither partially nor fully retired, that is, $w^{FT}_{it} = w^{FT^*}_{it} * (1 - d^R_{it})$.

The selection equation is estimated by probit regression, and the results are used to generate the inverse Mill's ratio, $\hat{\lambda}^R_{it}$. In the second stage, $\hat{\lambda}^R_{it}$ is added as a regressor to the wage equation, which is estimated by fixed-effects using observations for full-time workers only. Wooldridge (1995) shows that a valid test of the null hypothesis of no selection is a t-statistic on the coefficient of $\hat{\lambda}^R_{it}$.

In line with the retirement literature that characterizes partial and full retirement as two separate statuses (Honig and Hanoch (1985a)), I also consider an alternative specification for the selection process that allows for observables to affect the two decisions differentially. In this case, selection is based on two indices, d^{PR} and d^{FR}, which equal 1 if the individual is partially and fully retired, respectively. Full-time wages are observed when both indicators are equal to 0. The specifications for the selection indices are analogous to (5), with the residuals in the two probit equations assumed jointly normal conditional on Z_i. The first stage is estimated by multinomial probit. Under the assumption of independent idiosyncratic error terms in the partial and full retirement selection equations, the two inverse Mill's ratios, $\hat{\lambda}^{PR}_{it}$ and $\hat{\lambda}^{FR}_{it}$, enter additively the wage equation and their coefficients are used to perform the selection test.

First-stage results.

Results from the first-stage regressions are shown on table 3. The first 2 columns show estimates for the binomial probit specification, and the next 4 for the multinomial probit

[17]The vector $\bar{Z}_i$ is 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 individual i is observed in the panel. As discussed in Wooldridge (1995), equation (5) can be viewed as the reduced form of a selection equation with an unobserved, time-constant individual effect which is specified as a correlated random effect (as in Chamberlain (1980)). The inclusion of $\bar{Z}_i$ is intended to capture its correlation with observables.

[18]No further functional form assumption is required on the joint distribution of ε_{it} and v_{it} in this context where the aim is to test (and not to correct) for sample selection.

one. For ease of interpretation, marginal effects are reported next to the coefficients for each variable. As has been done before in the literature, the social security early and full retirement ages (ERA and FRA) are used as exclusion restrictions.[19] At the ERA and FRA, individuals face strong incentives to reduce their labor supply (see Aaronson and French (2004) for a detailed discussion). At the same time, these exogenous policy variables can be plausibly excluded from the second stage as long as employers are not systematically lowering wages of employees who become eligible for social security benefits. To my knowledge, there is no evidence that this is the case. In order to use the 62 and 65 age dummies as exclusion restrictions, I make the additional assumption that the participation and wage profiles outside these ages are well approximated by a smooth function of age.

The estimates in table 3 show that the social security ERA and FRA are statistically significant in all specifications and have the expected sign. Specification (1) shows that full-time participation declines by 14 percentage points at the ERA. The results from specification (2) indicate that up to four-fifths of this decline are attributable to transitions into full retirement, and the rest to transitions into part-time work. Instead, the decline in full-time participation at age 65 is mostly explained by transitions into partial retirement.

The rest of the regressors are divided in two groups, namely time-variant and time-invariant ones. The latter, which include time averages of time varying regressors together with variables that are constant in time, such as education, are intended in this framework to capture time-constant, individual effects. Focusing on time-invariant variables first, we can see that average age has a significant, negative impact on both partial and full retirement. Given the design of the the HRS panel, later-born workers enter the sample at younger ages and hence have a lower average age. This result therefore confirms Cahill, Giandrea and Quinn's (2006) finding that later-born workers are more likely to take on bridge jobs, and suggests that the increase in older men's labor force participation over the last two decades (Schirle (2008)) may be partly attributable to increases in part-time participation.

Not surprisingly, the lower an individual's health type –proxied by the average number of periods in which he reports suffering from a health condition that limits his ability to work– the lower the likelihood of full-time work. A one-standard deviation increase in the average health limitation variable is associated with a 2.2% increase in the probability of part-time work and a 5.5% increase in the probability of retirement. Lifetime wealth, measured here

[19] See, for example, Aaronson and French (2004) and Bernheim, Skinner and Weinberg (2001). The ERA is 62 for all individuals in the sample, while the FRA varies from 65 years to 65 years and 8 months. I make no distinction between those who reach the FRA at the beginning or closer to the end of their 65^{th} year.

by the average of wealth holdings over the period, is not significant in the partial retirement equation, but it is positively associated with full retirement: A one standard deviation increase in lifetime wealth increases the full retirement probability by 4.1%.

College-educated individuals are more likely than those without a high school degree (the omitted category) to work part time, and less likely to fully retire. The corresponding marginal effects for high school graduates go in the same direction, but their magnitude is smaller and they are only significant at 10%. The type of pension plan an individual holds is also a strong predictor of retirement. Workers who have a defined benefit plan in their full-time job are 6 percentage points less likely to partially retire, and slightly more likely to fully retire than those who have no pension (the omitted category). Those with a defined contribution plan in their full-time job are the most likely to continue working full time. Mother's education, which aims to proxy for individual ability, is not associated with retirement propensities.

Turning now to the time-variant regressors, we can see that the probability of both partial and full retirement increases by around 4 percentage points every year. The onset of a health limitation decreases the probability of partial retirement by 4 percentage points, while increasing the probability of full retirement by more than 20 percentage points. Finally, transitory changes in wealth have essentially no effect on retirement transitions.

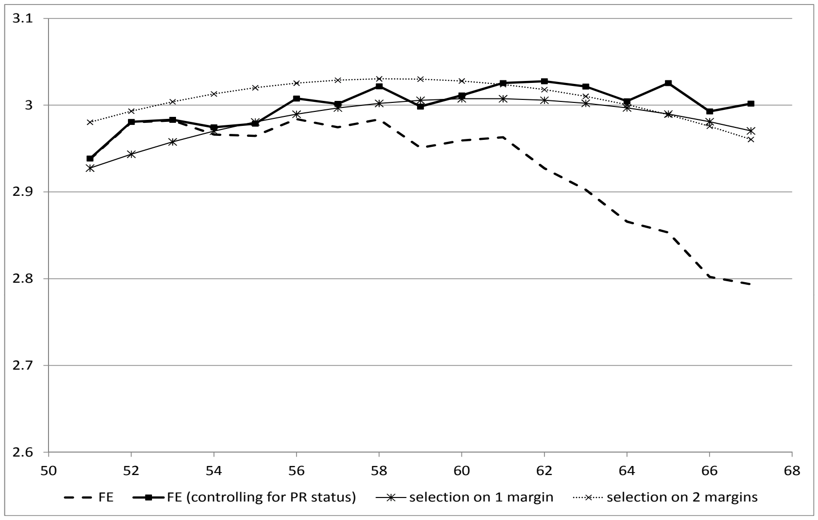
The conclusion from the first stage results is that retirement decisions are strongly associated with observables. The fact that these same observables correlate with wage levels generates composition effects that lead to the differences between the OLS and fixed-effects wage profiles discussed in section 3.

Second-stage results.

Results for the second stage are presented in table 4. The two columns correspond to the specifications with one and two margins of selection, respectively. The coefficient on the inverse Mill's ratio in the first column indicates positive selection, that is, workers who receive positive idiosyncratic shocks to wages are less likely to partially or fully retired. The coefficients in the second equation indicate positive selection into partial retirement, but negative selection into full retirement. The magnitude of all three coefficients is relatively small, and neither is statistically significant. In the second column, the p-value for the joint significance test of the two inverse Mill's ratios is 0.61. Thus the null hypothesis of no selection bias cannot be rejected for any specification.

Despite the statistical rejection, it is worth considering whether the magnitude of the coefficients on the inverse Mill's ratios would be negligible in practice. This in important in

Figure 8: Log Wage Profiles from Second Stage Regression Estimates



NOTE. - "F.E". and "FE(controlling for PR status)" series constructed using age coefficients from columns 2 and 3 in table 1. "Selection on 1 margin" and "Selection on 2 margins" constructed using age coefficients from columns 1 and 2 in table 4.

order to ascertain that the rejection of self-selection is not simply due to a lack of precision in the estimates for those coefficients. In the presence of selection, the inverse Mill's ratio is an omitted variable in the regressions estimated in section 3. To the extent that the omitted variable is correlated with the age coefficients, we would expect to observe changes in the estimates of the age profile once the inverse Mill's ratio is added to regression. Table 4 shows that the age coefficients are significant in the two specifications, indicating a concave profile. However, a graphical analysis of the data (see figure 8) shows that the concavity results from fitting the initial increase and subsequent flattening of hourly wages with a quadratic polynomial. Indeed, the age-wage profiles from the two second-stage specifications remain very close to the one estimated in section 3 controlling for transitions into partial retirement. Hence there is no evidence that the flat full-time wage profile suffers from omitted variable bias arising from self-selection into retirement as a result of wage shocks.[20]

The rejection of the self-selection model can be interpreted in two ways. First, it is possible that condition (3) is not violated because individuals do not select into retirement on the basis of realizations of u_{it}. The first-stage results underline the strong role of observables in driving retirement decisions, and the combined effect of pension type, health realizations,

[20]It is important to keep in mind that the procedure followed in this section allows to test, but not to correct, for selection. Given that selection on unobservables is rejected, the preferred estimate for the age-wage profile is still the one presented in column 3 of table 1, which uses a less restrictive age specification.

education, wealth, etc. may override any potential role of idiosyncratic shocks in work exits. Second, the variance of transitory shocks could be too small to give rise to retirement transitions. Idiosyncratic shocks may trigger small hour adjustments but still be below too small to explain the large, discrete hour changes observed for workers who partially or fully retire. In either case, the evidence presented above indicates that self-selection is not an important driver of retirement transitions. A straightforward implication of the absence of selection is that the age profile of offered wages is nondecreasing in age.

B. Voluntary vs. Involuntary Retirement

Having established that retirement transitions are not driven by declines in offered wages, I now address the question of whether they are voluntary. In other words, are transitions out of full-time work driven by a relative preference for leisure or by exogenous employment constraints? Distinguishing between the two is not straightforward, given that retirement transitions are influenced by a myriad of motives and that the definition of "voluntary" retirement is somehow subjective. Consider, for example, two workers who are diagnosed with arthritis. One of them may view his retirement as involuntary, the result of an unexpected health condition that makes work too painful. The second one, who has higher savings and access to a pension plan, may claim that he voluntarily retired to avoid the rising cost of work. The aim of this section is not to provide a formal definition of what constitutes a "voluntary" retirement. Rather, I review several pieces of evidence from different sources that, on the whole, point to a majority of retirements being a choice made by workers who had the option but chose not to remain in their full-time jobs.

The first sources of evidence are papers showing that most workers plan to either partially or fully retire at some point in the future (Abraham and Houseman (2004)) and are able to predict their retirement date with high accuracy, even when asked years in advance (Bernheim (1989), Haider and Stephens (2004)). These studies indicate that most retirement transitions are, at least, expected. While this is reassuring, it is not enough to conclude that they are also voluntary.

More direct evidence comes from self reports obtained from surveys of individuals who have already retired. Szinovacz and Davey (2005) find that, when asked directly, 70% of male HRS retirees perceive their retirement as "wanted" rather than "forced." In interpreting these results, we may be concerned that responses to this type of question could be contaminated by justification bias if workers are reluctant to admit that they were coerced out of their job. I complement Szinovacz and Davey's evidence by analyzing a series of questions in the HRS survey regarding the reasons that lead workers to leave a job. These questions are asked of all HRS respondents who change employer between two waves, regardless of retirement status. Since workers are not asked to give a subjective evaluation of the reason why they left their previous job, their answers to these questions are arguably less susceptible to justification bias.

Results are presented in table 5. The first two columns show the reasons for leaving their last employer cited by workers who fully and partially retired between the last two waves, respectively. For comparison, the third and fourth columns show the answers of workers who moved between full-time jobs (i.e., they changed employer but did not retire) and those who were already retired in the first wave (most of these workers are transiting from part-time jobs into full retirement). I group the reasons for leaving an employer according to how likely they are to indicate an involuntary job separation. Focusing first on involuntary transitions, the comparison of columns 1 and 2 shows that full and partial retirees are almost equally likely to have left their last job due to employment constraints (i.e., being laid off or displaced because of business closure) or to negative changes in their working conditions. However, full retirees are four time as likely as partial retirees to have been forced to retire because of poor health or a disability, consistent with the results from the first stage regression estimation in the previous section. Overall, one-third of full retirees and one-fifth of partial retirees report having left their previous full time job for involuntary reasons. The remaining job separations are due to voluntary reasons. Most of these (87% of full retirees and 70% of partial retirees) are attributed to a desire to enjoy more leisure.

By comparison, full-time workers who switch employer but do not retire (column 3) are much more likely than partial and full retirees to have suffered exogenous changes in their working conditions and less likely to have had health problems. Finally, transitions from partial into full retirement (column 4) seem to respond to very similar motives as transitions from full-time work into full retirement.

Even though not shown separately in the table, one of the possible answers to the question asking for reasons leading to job transitions is that wages would have been cut had the worker stayed with the previous employer. It is worth pointing out that only 1 out 672 workers who report having entered full retirement voluntarily, and none of those who transited voluntarily into partial retirement, mention this reason for their work transition.

The evidence presented in this section confirms Szinovacz and Davey's (2005) results and indicates that the largest share of full-time work exits is voluntary. The number of involuntary retirements, between 20 and 30% of the total, is far from negligible, and points to the key role of uncertainty in the labor supply decisions of older workers. However, up to 70% of full retirements, and 80% of partial retirements, arise as the optimal choice for workers who could have remained employed full time at their previous wage. For all those workers, the utility associated to increasing leisure outweighs that associated to full-time employment.

The conclusions from this section are twofold. First, the offered full-time wage profile is nondecreasing in age. Second, hours choices, and in particular retirement transitions, are not a response to declining wages. Rather, hours and *ex-post* wages are determined simultaneously for a majority of workers.

5 Conclusions

The large literature studying the wage process has paid careful attention to wage uncertainty, but comparatively less focus has been placed on the deterministic component of wages, particularly as workers approach the end of their careers. However, key results in the labor supply literature rely on the correct specification of the wage profile. These include the measurement of income uncertainty, the estimation of the intertemporal elasticity of substitution of labor supply, or the assessment of responses to retirement incentives.

This paper shows that the offered hourly wage profile does not decline at older ages for as long as workers remain employed full-time. Wage declines set in only at the time of transition from full-time into part-time work. The hump-shaped profile often estimated in the literature results from pooling observations of full- and part-time workers, and is not a good proxy for the wage path followed by the typical individual.

The paper also shows that the flat wage profile observed in the data is not driven by selection into retirement of workers who receive negative wage shocks. Finally, for the majority of workers wage declines are determined endogenously. These workers have the opportunity to remain employed full time at their previous wage, but choose instead to enjoy more leisure and work part time for a lower wage. In order to accurately reflect the typical worker's decision process, retirement models must allow for wages and hours to be chosen simultaneously, rather than represent retirement as the optimal response to an exogenously declining wage profile.

References

[1] Aaronson, D. and E. French. 2004. "The Effect of Part-Time Work of Wages: Evidence from the Social Security Rules". *Journal of Labor Economics*. Vol. 22, pp. 329-352.

[2] Abraham, K. and S. Houseman. 2005. "Work and retirement plans among older Americans", in R. Clark and O. Mitchell (eds.), *Reinventing the Retirement Paradigm*, Oxford: Oxford University Press.

[3] Becker, G. S. 1993. *Human Capital: A Theoretical and Empirical Analysis, with Special Reference to Education. 3rd Edition.* The University of Chicago Press.

[4] Ben-Porath, Y. 1967. "The Production of Human Capital and the Life-Cycle of Earnings". *Journal of Political Economy*. Vol. 75, pp. 352-365.

[5] Bernheim, B. D. 1989. "The Timing of Retirement: A Comparison of Expectations and Realizations", in D. A. Wise (ed.), *The Economics of Aging*, Chicago and London: Unversity of Chicago Press.

[6] Bernheim, B. D., J. Skinner and S. Weinberg. 2001. "What Accounts for the Variation in Retirement Wealth among U.S. Households?" *The American Economic Review*. Vol. 91(4), pp. 832-857

[7] Blau, D. 2008. "Retirement and Consumption in a Life Cycle Model". *Journal of Labor Economics*. Vol. 26(1), pp. 35-71.

[8] Cahill, K. E., M. D. Giandrea and J. F. Quinn. 2006. "Retirement Patterns From Career Employment". *The Gerontologist*. Vol. 46(4), pp. 514-523

[9] Casanova, M. 2013. "Misspecified Wage Profiles and Estimates of Labor Supply Elasticities of Older Workers". *Mimeo* UCLA.

[10] Chamberlain, G. 1980. "Analysis of Covariance with Qualitative Data". *Review of Economic Studies*. Vol. 47, pp. 225-238.

[11] Flodén, M. and J. Lindé. 2001. "Idiosyncratic Risk in the United States and Sweden: Is There a Role for Government Insurance?". *Review of Economic Dynamics*. Vol. 4(2), pp. 406-437.

[12] French, E. 2005. "The Effects of Health, Wealth, and Wages on Labour Supply and Retirement Behaviour". *Review of Economic Studies*. Vol. 72, pp. 395-427

[13] Ghez, G. and G. S. Becker. 1975. "The Allocation of Goods Over the Life Cycle". In *The Allocation of Time and Goods over the Life Cycle*. New York: National Bureau of Economic Research, Columbia University Press.

[14] Gomes, F. J., L. J. Kotlikoff, L. M. Viceira. 2008. "Optimal Life-Cycle Investing with Flexible Labor Supply: A Welfare Analysis of Life-cycle Funds". *The American Economic Review*. Vol. 98 (2), pp. 291-303. Papers and Proceedings of the One Hundred Twentieth Annual Meeting of the American Economic Association.

[15] Gourinchas, P. O. and J. A. Parker. 2002. "Consumption Over the Life Cycle". *Econometrica*.

Vol. 70(1), pp. 47-89.

[16] Gustman, A. and T. Steinmeier. 1985. "The Effect of Partial Retirement on Wage Profiles of Older Workers". *Industrial Relations.* Vol. 24, pp. 257-265.

[17] Guvenen, F. 2009. "An empirical investigation of labor income processes". *Review of Economic Dynamics.* Vol. 12(1), pp. 5879

[18] Haider, S. J. and M. Stephens Jr. 2007. "Is there a retirementconsumption puzzle? Evidence using subjective retirement expectations". *Review of Economics and Statistics.* Vol. 84, pp. 24764.

[19] Hall, R. E. and F. S. Mishkin. "The Sensitivity of Consumption to Transitory Income: Estimates from Panel Data on Households". *Econometrica.* 50 (1982) 461-481

[20] Heathcote, J., K. Storesletten and G. L. Violante. 2010. "The Macroeconomic Implications of Rising Wage Inequality in the United States". *Journal of Political Economy.* Vol. 118(4), pp. 681-722.

[21] Heckman, J. J. 1976. "A Life-Cycle Model of Earnings, Learning, and Consumption". *Journal of Political Economy.* Vol. 84(4), pp.S11-S44.

[22] Heckman, J. J. 1979. "Sample selection bias as a specification error". *Econometrica.* Vol. 47, pp. 153-61.

[23] Heckman, J. J., L. J. Lochner and P. Todd. 2006. "Fifty Years of Mincer Earnings Regrressions". In Hanushek, E. and F. Welch, *Handbook of Education Economics. Vol. 1.* (The Netherlands: Elsevier)

[24] Honig, M. and G. Hanoch. 1985a. "Partial Retirement as a Separate Mode of Retirement Behavior". *The Journal of Human Resources.* Vol. 20, pp. 21-46.

[25] Honig, M. and G. Hanoch. 1985b. "'True' Age Profiles of Earnings: Adjusting for Censoring and for period and Cohort Effects". *The Review of Economics and Statistics.* Vol. 67, pp. 383-394.

[26] Hubbard, R. G., J. Skinner and S. P. Zeldes "Precautionary Saving and Social Insurance". *Journal of Political Economy.* 103 (1995) 360-399

[27] Hurd, M. 1971. "Changes in Wage Rates Between 1959 and 1967, *The Review of Economics and Statistics.* Vol. 53, pp. 189-199.

[28] Hurd, M. 1996. "The Effect of Labor Market Rigidities on the Labor Force Behavior of Older Workers" in D. A. Wise (ed.), *Advances in the Economics of Aging*, The University of Chicago Press

[29] Johnson, R. W. and D. Neumark. 1996. "Wage Declines among Older Men. *The Review of Economics and Statistics.* Vol. 78, pp. 740-748.

[30] Kopecky, K. A. 2011. "The Trend in Retirement". *International Economic Review.* Vol. 52(2), pp. 287316

[31] Krueger, D. and F. Kubler. 2006. “Pareto-Improving Social Security Reform when Financial Markets Are Incomplete?”. *The American Economic Review*. Vol. 96(3). pp. 737-755.

[32] Lazear, E. “Why Is There Mandatory Retirement?” *Journal of Political Economy*. 87 (1979) 1261-1284.

[33] Lillard, L. A. and R. J. Willis. 1978. “Dynamic Aspects of Earning Mobility”. *Econometrica*. Vol. 46, pp. 985-1012.

[34] Low, H. W. 2005. “Self-insurance in a life-cycle model of labour supply and savings”. *Review of Economic Dynamics*. Vol. 8(4), pp. 945-975.

[35] Maestas, N. and J. Zissimopoulos. 2010. “How Longer Work Lives Ease the Crunch of Population Aging”. *Journal of Economic Perspectives*. Vol. 21, pp. 139-160.

[36] Mincer, J. 1974. *Schooling, Experience, and Earnings*. New York: Columbia University Press.

[37] Murphy, K. and F. Welch. 1990. “Empirical Age-Earnings Profiles”. *Journal of Labor Economics*. Vol. 8(2), pp. 202-229

[38] Murphy, K. and F. Welch. 1992. “The Structure of Wages”. *Quarterly Journal of Economics*. Vol. 107(1), pp. 285-326.

[39] Ruggles, S., J. T. Alexander, K. Genadek, R. Goeken, M. B. Schroeder, and M. Sobek. 2010. “Integrated Public Use Microdata Series: Version 5.0 [Machine-readable database]”. Minneapolis: University of Minnesota.

[40] Ruhm, C. J. 1990. “Bridge Jobs and Partial Retirement”. *Journal of Labor Economics*. Vol. 8, pp. 482-501.

[41] Rupert, P. and G. Zanella. 2012. “Revisiting Wage, Earnings and Hours Profiles”. UCSB Working Paper.

[42] Rust J. and C. Phelan. 1997. “How Social Security and Medicare Affect Retirement Behavior in a World of Incomplete Markets”. *Econometrica*. Vol. 65, pp. 781831.

[43] Scholz, J. K., A, Seshadri and S. Khitatrakun. 2006. “Are Americans Saving Optimally for Retirement?”. *Journal of Political Economy*. Vol. 114(4), pp. 607-643.

[44] Schirle, T. 2008. “Why Have the Labor Force participation Rates of Older Men Increased since the Mid-1990s?”. *Journal of Labor Economics*. Vol. 26(4), pp. 549-594.

[45] St.Clair, P., D. Blake, D. Bugliari, S. Chien, O. Hayden, M. Hurd, S. Ilchuk, F. Kung, A. Miu, C. Panis, P. Pantoja, A. Rastegar, S. Rohwedder, E. Roth, J. Carroll, J. Zissimopoulos. 2010. “RAND HRS data documentation: Version J”. Santa Monica, CA: RAND Center for the Study of Aging.

[46] Storesletten, K., C. I. Telmer and A. Yaron. 2004. “Cyclical Dynamics in Idiosyncratic Labor Market Risk”. *Journal of Political Economy*. Vol. 112(3), pp. 695-717

[47] Szinovacz, M. E. and A. Davey. 2005. “Predictors of Perceptions of Involuntary Retirement”. *The Gerontologist*. Vol. 45, pp. 36-47.

[48] Welch, F. 1979. “Effects of Cohort Size on Earnings: The Baby Boom's Babies Financial Bust”. *Journal of Political Economy.* Vol. 87, pp. 65-98.

[49] Willis, R. J. 1985. “Wage Determinants: A Survey and Reinterpretation of Human Capital Earnings Functions. In O. Ashenfelter and D. Card (eds.), *Handbook of Labor Economics.* North Holland, Amsterdam.

[50] Wooldridge, J. M. 1995. “Selection Correction for Panel Data Models Under Conditional Mean Independence Assumptions”. *Journal of Econometrics.* Vol. 68, pp. 115-132.

Table 1: Age-Log Wage Profiles for Different Specifications

	Dependent Variable				
	w_{it}	w_{it}	w_{it}	w_{it}^{FT}	w_{it}^{PT}
	OLS	FE	FE	FE	FE
	(1)	(2)	(3)	(4)	(5)
PR=1			-0.337***		
			(0.025)		
age≥52	0.021	0.042	0.042	0.035	
	(0.039)	(0.030)	(0.029)	(0.023)	
age≥53	-0.040	0.002	0.002	0.003	
	(0.034)	(0.027)	(0.025)	(0.019)	
age≥54	0.018	-0.016	-0.009	0.011	
	(0.030)	(0.022)	(0.020)	(0.016)	
age≥55	-0.003	-0.002	0.004	0.007	
	(0.029)	(0.020)	(0.019)	(0.015)	
age≥56	0.011	0.019	0.029*	0.021	
	(0.025)	(0.017)	(0.016)	(0.012)	
age≥57	-0.013	-0.009	-0.006	0.010	
	(0.024)	(0.016)	(0.015)	(0.012)	
age≥58	-0.007	0.009	0.020	-0.003	0.072
	(0.024)	(0.015)	(0.015)	(0.012)	(0.074)
age≥59	-0.019	-0.033**	-0.023	0.003	-0.087
	(0.024)	(0.016)	(0.015)	(0.012)	(0.063)
age≥60	0.002	0.008	0.013	0.012	-0.124**
	(0.024)	(0.017)	(0.015)	(0.012)	(0.060)
age≥61	-0.019	0.004	0.014	0.001	0.136**
	(0.024)	(0.016)	(0.016)	(0.013)	(0.062)
age≥62	-0.044	-0.036**	0.002	-0.002	0.002
	(0.028)	(0.018)	(0.017)	(0.015)	(0.049)
age≥63	-0.032	-0.025	-0.006	0.014	0.002
	(0.033)	(0.021)	(0.020)	(0.017)	(0.041)
age≥64	-0.080**	-0.037	-0.017	-0.024	-0.007
	(0.036)	(0.022)	(0.021)	(0.018)	(0.040)
age≥65	-0.006	-0.013	0.021	0.014	0.072*
	(0.041)	(0.026)	(0.024)	(0.023)	(0.037)
age≥66	-0.082*	-0.051	-0.033	-0.007	-0.076*
	(0.047)	(0.033)	(0.031)	(0.029)	(0.041)
age≥67	0.014	-0.008	0.009	-0.044	0.049
	(0.048)	(0.035)	(0.033)	(0.035)	(0.041)
individual-year obs.	7,915	7,500	7,500	6,277	830
# of individuals		1,834	1,834	1,666	287
R^2	0.22	0.05	0.15	0.01	0.02
Tests of Joint Significance (p-value):					
Age≥52-Age≥60	0.695	0.059	0.08	0.012	
Age≥61-Age≥67	0.000	0.000	0.618	0.733	0.01

NOTE. - Robust standard errors in parentheses. *, ** and *** indicate significance at 10, 5, and 1%, respectively. All regressions include an intercept and a measure of the unemployment rate to capture yearly changes in labor market conditions. 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-Log Hours Profiles for Different Specifications

	Dependent Variable				
	$hours_{it}$	$hours_{it}$	$hours_{it}$	$hours_{it}^{FT}$	$hours_{it}^{PT}$
	OLS	FE	FE	FE	FE
	(1)	(2)	(3)	(4)	(5)
PR=1			-0.642***		
			(0.025)		
age≥52	-0.032**	-0.030	-0.030	-0.013	
	(0.016)	(0.022)	(0.021)	(0.013)	
age≥53	0.024*	0.038*	0.038**	0.002	
	(0.013)	(0.020)	(0.019)	(0.010)	
age≥54	-0.002	-0.024	-0.010	0.002	
	(0.011)	(0.018)	(0.015)	(0.009)	
age≥55	-0.006	-0.000	0.011	-0.007	
	(0.012)	(0.017)	(0.014)	(0.009)	
age≥56	-0.016	-0.027*	-0.009	-0.013*	
	(0.011)	(0.016)	(0.013)	(0.008)	
age≥57	-0.007	0.009	0.015	-0.001	
	(0.011)	(0.016)	(0.013)	(0.007)	
age≥58	-0.007	-0.019	0.002	-0.008	0.068
	(0.013)	(0.017)	(0.014)	(0.008)	(0.075)
age≥59	-0.012	-0.002	0.016	0.011	-0.042
	(0.014)	(0.018)	(0.015)	(0.008)	(0.081)
age≥60	-0.008	-0.026	-0.018	-0.015**	0.037
	(0.015)	(0.017)	(0.014)	(0.007)	(0.064)
age≥61	0.004	-0.003	0.017	0.004	-0.017
	(0.014)	(0.017)	(0.014)	(0.007)	(0.066)
age≥62	-0.105***	-0.106***	-0.034**	-0.021***	-0.049
	(0.021)	(0.021)	(0.016)	(0.008)	(0.061)
age≥63	-0.055*	-0.065**	-0.030	0.002	-0.109
	(0.030)	(0.029)	(0.022)	(0.009)	(0.063)
age≥64	-0.048	-0.034	0.004	-0.002	0.045
	(0.037)	(0.033)	(0.026)	(0.011)	(0.061)
age≥65	-0.048	-0.060*	0.004	-0.013	0.004
	(0.041)	(0.035)	(0.029)	(0.013)	(0.057)
age≥66	-0.032	-0.027	0.008	-0.008	0.005
	(0.047)	(0.040)	(0.034)	(0.016)	(0.054)
age≥67	-0.012	-0.015	0.019	0.011	0.077
	(0.050)	(0.043)	(0.037)	(0.018)	(0.056)
individual-year obs.	7,915	7,500	7,500	6,277	830
# of individuals		1,834	1,834	1,666	287
R^2	0.10	0.15	0.42	0.02	0.02
Tests of Joint Significance (p-value):					
Age≥52-Age≥60	0.000	0.021	0.341	0.043	
Age≥61-Age≥67	0.000	0.000	0.000	0.001	0.01

NOTE. - Robust standard errors in parentheses. *, ** and *** indicate significance at 10, 5, and 1%, respectively. All regressions include an intercept and a measure of the unemployment rate to capture yearly changes in labor market conditions. 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 binomial (1) and multinomial (2) probit estimates.

	(1)		(2)			
	PT+Fully Ret.		PT		Fully Ret.	
	Coeff	Mg Effects	Coeff	Mg Effects	Coeff	Mg Effects
Time-Varying Regressors						
age	-0.118	0.040***	0.167	0.035***	-0.271	0.042***
	(0.191)	(0.002)	(0.271)	(0.004)	(0.233)	(0.003)
age^2	0.002		0.000		0.004**	
	(0.002)		(0.002)		(0.002)	
health limit	0.672***	0.165***	0.252***	-0.043***	1.107***	0.206***
	(0.052)	(0.011)	(0.084)	(0.008)	(0.068)	(0.013)
current wealth (in 100k)	0.022	0.005	-0.001	0.002	0.044	0.007*
	(0.021)	(0.003)	(0.032)	(0.004)	(0.030)	(0.004)
current wealth^2	-0.000		0.002		-0.001	
	(0.001)		(0.002)		(0.002)	
Exclusion Restrictions						
over ERA	0.531***	0.143***	0.564***	0.025**	0.768***	0.119***
	(0.060)	(0.015)	(0.082)	(0.010)	(0.074)	(0.015)
over FRA	0.150*	0.036**	0.263***	0.022**	0.168*	0.014
	(0.079)	(0.016)	(0.097)	(0.011)	(0.087)	(0.015)
Time-Constant Regressors						
avg. age	1.185***	-0.024***	1.578**	-0.027***	1.544***	-0.023***
	(0.270)	(0.003)	(0.617)	(0.004)	(0.538)	(0.003)
avg. age^2	-0.011***		-0.014***		-0.014***	
	(0.002)		(0.005)		(0.004)	
avg. health limit	0.675***	0.158***	0.427**	0.077**	1.090***	0.189***
	(0.077)	(0.024)	(0.182)	(0.033)	(0.145)	(0.025)
avg. wealth (in 100k)	0.073***	0.011**	0.044	0.004	0.122***	0.015***
	(0.024)	(0.005)	(0.053)	(0.006)	(0.043)	(0.005)
avg. wealth^2	-0.004**		-0.004		-0.006*	
	(0.002)		(0.004)		(0.003)	
education = college	-0.104**	-0.024	0.113	0.030**	-0.254**	-0.053***
	(0.043)	(0.016)	(0.120)	(0.013)	(0.102)	(0.016)
education = high school	-0.028	-0.007	0.113	0.021*	-0.109	-0.027*
	(0.040)	(0.015)	(0.111)	(0.012)	(0.091)	(0.015)
mother's education	-0.014	-0.003	0.046	0.008	-0.050	-0.009
	(0.015)	(0.006)	(0.042)	(0.008)	(0.035)	(0.006)
pension type = DB	-0.143***	-0.034*	-0.532***	-0.064***	-0.023	0.030*
	(0.045)	(0.019)	(0.123)	(0.013)	(0.106)	(0.016)
pension type = DC	-0.381***	-0.089***	-0.532***	-0.032***	-0.493***	-0.057***
	(0.048)	(0.018)	(0.132)	(0.012)	(0.114)	(0.017)
Obs.	11,877		11,877		11,877	

NOTE. - *, ** and ** indicate significance at 10, 5, and 1%, respectively. Robust standard errors reported for regression coefficients. Standard errors for marginal effects are computed using 1,000 bootstrap replications clustered at the individual level in order to account for the panel structure of the data. All regressions include an intercept and a measure of the unemployment rate to capture yearly changes in labor market conditions. Observations with wealth above $2M (1.31% of the sample) are dropped from specifications where wealth appears as regressor.

Table 4: Second-stage estimates.

	(1)	(2)
age	0.107***	0.109***
	(0.032)	(0.036)
age^2	-0.001***	-0.001***
	(0.000)	(0.000)
health limit	-0.031*	-0.016
	(0.017)	(0.022)
wealth (in 100k)	0.005	0.006
	(0.005)	(0.005)
wealth^2	-0.000	-0.000
	(0.000)	(0.000)
$\hat{\lambda}^R$	-0.012	
	(0.029)	
$\hat{\lambda}^{PR}$		-0.057
		(0.057)
$\hat{\lambda}^{FR}$		0.028
		(0.035)
individual-year observations	6,244	6,244
# of individuals	1,661	1,661

NOTE. - *, ** and ** indicate significance at 10, 5, and 1%, respectively. Standard errors (in parentheses) computed 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 Mill's ratios in the second stage. All regressions include an intercept and a measure of the unemployment rate to capture yearly changes in labor market conditions. Observations with wealth above $2M (1.31% of the sample) are dropped from specifications where wealth appears as regressor.

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 in t	
Business closed	4.20	3.97	18.30	6.31
Laid off/let go	9.13	10.60	24.55	13.46
Retired, negative change in working conditions	2.92	2.65	0.67	2.33
Poor health/disabled	17.26	4.64	2.23	15.45
Total involuntary retirements	**33.51**	**21.86**		
Leisure	53.15	50.33	11.16	36.05
Voluntary quits	3.20	15.89	34.82	14.62
Financial incentive	4.11	4.64	1.56	1.83
Change in external circumstances	0.91	0.66	2.01	2.16
Total voluntary retirements	**61.37**	**71.52**		
Other	5.11	6.62	4.69	7.81
Obs.	1,095	151	448	602

Source: author's calculations combining answers to the questions "why did you leave [your previous] employer?" and "did your employment situation change in some way that encouraged you to leave?" The category "Retired, negative change in working conditions" includes workers who cite retirement as the reason they left their last employer and, when asked about the change in their working conditions, report that a) their supervisor or coworkers encouraged their departure, b) their wages or hours had or would have been cut if they had stayed, c) they would have been laid off, d) they had new job duties, e) they had to move to a new job location or f) their employer had changed their health insurance. The category "Leisure" includes workers who say they left their employer because they wanted to travel, or those who claim to have retired and report no changes in their working conditions. The category "Voluntary quits" includes workers who report having left their employer because they a) quit, b) found a better job, c) sold their business, or d) handed over responsibilities to other family members. The category "Financial incentive" includes workers who took an early retirement incentive, became eligible for a pension, found it financially advantageous not to work, or found a job with better pay or benefits. The category "Change in external circumstances" includes those who had to stop working to care for a family member, because their family moved, or because they separated or divorced. Finally, the category "Other" includes workers who do not give the reason why they left their previous employer, together with those who provide reasons that are difficult to classify as voluntary or involuntary, namely a) transportation, distance to work, b) respondent/spouse was transferred (but we don't observe which one) and c) didn't get along with coworkers.

NOTE: The proportions in the table are computed using only nonmissing responses to the relevant questions.

Table 6: Robustness checks

	Sample definition	coeff. on PR=1	Joint sig. tests (p-value) Ages 52-60	Ages 61-67	N (total)	N (# ind.)	# partial retirements
0	Baseline sample	-0.337***	0.08	0.62	7,500	1,834	524
Wage limits							
1	$w_{it} \leq \$50$	-0.347***	0.05	0.88	7,211	1,789	508
2	$w_{it} \leq \$100$	-0.333***	0.043	0.752	7,592	1,849	527
3	$w_{it} \leq \infty$	-0.303***	0.017	0.400	7,653	1,857	532
Definition of partial retirement							
4	Keep all old-age self-employed (DEF1)	-0.342***	0.025	0.677	8,915	2,155	769
5	Keep PT, old-age self-employed (DEF2)	-0.350***	0.005	0.455	8,436	2,051	665
6	Use tenure changes (DEF3)	-0.278***	0.002	0.146	7,500	1,834	752
Results by education category							
7	H.S. dropouts	-0.206***	0.819	0.942	1,661	431	104
8	H.S. graduates	-0.322***	0.387	0.700	2,788	683	205
9	(Some) College	-0.412***	0.527	0.196	3,038	716	213

NOTE. - *, ** and ** indicate significance at 10, 5, and 1%, respectively.

Table 7: Descriptive Statistics by Work Status in First Sample Wave. Non self-employed only.

	Fully Retired	PT Workers	FT Workers
Age	57.37 (4.26)	56.26 (4.10)	55.37 (3.27)
Education (%)			
Less than High School	35.60	22.17	23.18
High School Graduate	34.90	36.79	36.57
(Some) College	29.50	41.04	40.26
Health Status (%)			
Disabled	31.94	2.79	0.15
Bad Health	23.70	18.14	12.05
Good Health	44.36	79.07	87.80
Median Wealth by Health Status (in 2011$)			
Disabled	24,500		
Bad Health	41,000	36,500	67,300
Good Health	123,000	99,000	100,000
Observations	1,152	215	2,663

NOTE. - Wealth is measured at the household level.

Appendix

A.1 Analysis of the Sample of "Previously Retired" Workers

The wage change at the time of separation from the full-time job is not observed for workers who are already fully or partially retired by the time they enter the HRS panel –in what follows, I will refer to these workers as the "previously retired". Notice that the wage change information may not be recovered even if we were to make use of retrospective SSA earnings records, which are available for a subsample of HRS respondents. Consider a worker who switches from full-time to part-time work in year t. His SSA records would show a decrease in earnings for that year. However, because retrospective hour measures are not available, we would not be able to map the earnings drop into an hourly wage drop.

Table 7 analyzes the characteristics of workers entering the HRS panel for the first time, classified according to their participation status. Workers in the first column are fully retired, those in the second column are partially retired, and those in the third one are working full time. Observations from the first two columns make up the "previously retired" sample, and are excluded from the main analysis.

Not surprisingly, "previously retired" workers are on average older. There are other noticeable differences across groups. In general, partially retired workers and full-time workers have similar characteristics, whereas full retirees are somewhat different. The latter are less educated and considerably more likely to be disabled or in bad health. The correlation between health status and net worth is striking. The wealth holdings of disabled individuals are about a fifth of those of individuals in good health. Those in bad health hold between one- and two-thirds as much wealth as those in good health.

The table suggests that the "previously retired" can be divided in two broad categories: one with low wealth and in bad health or disabled, and another with high wealth and good health. We can speculate that workers in the first category are more likely to have left their full-time job due to their health conditions, while those in the second one are more likely to have done so because they had the financial means to afford more leisure in the form of early retirement. It is worth pointing out that these two motives for retirement are present in the sample used in the main analysis. In particular, 50% of those respondents who become disability recipients before age 60 belong to the sample used in the analysis. So while the empirical sample undersamples excludes workers who retire very early in their career through disability insurance, disability benefit recipients are by no means excluded from the analysis.

A.2 Robustness Checks

Definition of partial retirement

Different studies have used different definitions of partial retirement. Maestas and Zissimopoulos (2010), for example, have pointed out that in some cases self-employment may be a form of partial retirement. In particular, workers who want to reduce hours and cannot do so at their current job, may opt for becoming self-employed. With this consideration in mind, I contemplate two alternative definitions of partial retirement. DEF1 classifies a worker as partially retired the first time he switches from full-time work into part-time work *or* self employment. DEF2 defines partial retirement as a transition from full-time work into part-time work or *part-time* self-employment. DEF0 is the definition used in the main analysis.

Notice that because of differences in the treatment of the self-employed, the 3 definitions impose different restrictions on the sample. DEF0 is the most restrictive. Any worker who reports being self-employed at any point during his career is dropped from the sample. Under DEF2, workers who enter the panel as full-time employees and transit into part-time self-employment before fully retiring are kept in the sample. DEF 1, the least restrictive, keeps observations of workers who enter the panel as full-time employees and transit into either full-time or part-time self-employment.

The 3 definitions also lead to differences in the prevalence of partial retirement. 29% of workers in the DEF0 sample are classified as partially retired in at least one period, versus 36% and 42% for the DEF2 and DEF1 samples, respectively.

Following Cahill, Giandrea and Quinn (2006), I also consider an alternative definition of partial retirement (DEF3) that makes use of job tenure information. In their paper, they consider that a worker becomes partially retired whenever he moves to a new job that he will keep for less than 10 years before fully retiring. This definition imposes stringent data requirements, as it is necessary to follow workers for at least 10 years after their last work transition. In the HRS, this is not possible in most cases. Under DEF3, I assume that all workers who switch jobs in their 50s become partially retired. Given that few of those who switch jobs in their 50s will end up accumulating more than 10 years of tenure in the new job before they retire, DEF3 should be a good approximation to the definition of partial used by Cahill et al. (2006). According to DEF3, 41% of workers become partially retired at some point in their career

As a robustness check, I have re-run the main regressions in the paper using definitions 1 to 3. Results are shown in rows 4 to 6 of table 6. For DEF1 and DEF2, the key results remain virtually unchanged, indicating that the conclusions regarding the wage profile are not sensitive to the treatment of the self-employed. For DEF3, the wage drop associated to entry into partial retirement is several percentage points smaller. Recall that this definition interprets all job-to-job transitions after age 50 as partial retirements. It is likely that some of these late-career job moves are undertaken by workers who are not entering retirement, which explains the dampening of the partial retirement coefficient. All other results regarding the shape of the wage profile stand unchanged; in particular, the absence of wage drops during the retirement years cannot be rejected.