HOURS FLEXIBILITY AND RETIREMENT

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Data from the Health and Retirement Study indicate that hours constraints are a common feature of jobs held by workers nearing retirement. We present a simple model that predicts that workers who are not free to lower their usual hours of work should be more likely than their unconstrained counterparts to retire by some future date. Our estimates, which are robust to various specifications, support this prediction. The amount by which being hours constrained is estimated to raise retirement probabilities is nearly as large as the effect of being in relatively poor health, suggesting an economically significant effect. (*JEL* J26, J22, J14)

I. INTRODUCTION

Standard models of labor supply depict workers as choosing to work any positive amount of hours they wish at the competitive wage paid to persons of their type. Implicit in these models is the assumption that employers offer to workers, at a particular wage, a menu of infinitely many hours options from which workers may choose. However, this assumption seems fundamentally at odds with how the labor market functions. In particular, the distribution of weekly hours of work suggests that workers may be offered only a limited set of hours options by the employers.[1] This study examines how employer-imposed minimum hours of work requirements affect workers' retirement decisions.[2]

We present a simple theoretical framework, which shows that workers who cannot lower their hours of work in one period are more likely to retire by some subsequent date than workers not so constrained. The intuition is straightforward. Given his characteristics and the wage he commands, there is a particular hours level which a person *desires* to work in any period. Equivalently, between any two time periods, a worker initially working a given amount of hours desires that his hours *change* over time by a particular amount. For workers with jobs that feature flexible hours, the *actual* change in hours worked over time always perfectly matches the *desired* change. If the worker desires to withdraw from the labor force, then he does so. If he desires to work fewer positive hours in the later period,

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1. Many authors have shown that the distribution of actual hours worked in the week peaks dramatically at exactly 40 h, with very few prime-aged men working between 30 and 35 h, for example. This type of distribution is very unlikely to be the result of individual worker preferences.

2. Several economists have studied hours constraints in the labor market, though the effect on retirement has received very little attention. Ashenfelter (1980), Biddle and Zarkin (1989), Deaton and Muellbauer (1981), Ham (1982), and Dickens and Lundberg (1993) all assessed the impact of hours constraints on hours of work in the cross-section, while Ball (1990) uses panel data. Altonji and Paxson (1988) studied how the association between job changes and changes in preferences is affected by hours restrictions in the labor market. Kahn and Lang (1991) use the Canadian Labor Force Survey to investigate how actual hours deviate from desired hours.

ABBREVIATION

HRS: Health and Retirement Study

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he changes his labor supply to precisely the desired level.

For workers with the same desired change in labor supply but who face hours constraints, the only way to lower hours of work from one positive amount to another is to change jobs. Because such job changes are associated with transition costs, the worker chooses from among a set of less than ideal options. He can continue at the present job and be "overemployed." Alternatively, he can pay a transition cost and work the desired hours level in a new job, or he can withdraw completely from the labor force. Thus, unlike the worker whose hours are flexible, the hours-constrained worker retires both when he desires to work 0 h *and* when he wishes to work some positive amount less than he did previously but faces relatively high transition costs.

We tested the prediction that hours-constrained workers are more likely to retire over time using a sample of retirement-aged individuals from the Health and Retirement Study (HRS). We found that a large fraction of workers aged 55–64 yr cannot lower their usual weekly hours of work in 1992. These constraints are prevalent for both men and women, among people working different hours level, and across all types of jobs. Exploiting the panel feature of the data, and using a series of empirical tests, we found that hours-constrained workers are much more likely than their unconstrained peers to be retired by 1996. We document few differences in observed characteristics between constrained and unconstrained workers, suggesting that the results do not derive from differences in latent retirement propensities across workers or to unmeasured job characteristics. Moreover, we controlled carefully for demographic and job information. We also conducted some simple tests to assess the likelihood that the results derive from sorting by job type (constrained or unconstrained) or simple measurement error and found that our main estimates continue to obtain.

The study's emphasis that hours constraints make it difficult for workers to satisfy changing labor supply preferences without changing jobs is also the main focus of Altonji and Paxson (1992). These authors showed that in a sample of prime-aged women, events likely to generate large changes in labor supply preferences, such as the birth of a child or a child starting school, tend to be associated with large changes in hours worked only for persons who change jobs. Like nearly all the previous literature on hours constraints, their work is not about retirement and uses a model quite different from the one we present. Nevertheless, our work is similar in spirit to that of Altonji and Paxson.

More directly related to the present study is work by Gustman and Steinmeier (1983, 1984). Gustman and Steinmeier (1983) reviewed evidence from the Panel Study of Income Dynamics and a survey of about 300 persons conducted by the American Society for Personnel Administration and Bureau of National Affairs. These authors showed that many retirement-aged workers were indeed hours constrained in the time period studied. Gustman and Steinmeier (1984) found that reductions in the intensity of labor supply tend to occur with job changes. The result is very similar to the findings by Altonji and Paxson (1992) discussed above. However, it is only indirect evidence regarding whether hours constraints affect retirement in the manner we suggest since no data on the hours-constrained status of individual workers are used. We extend this previous work.

Our results are consistent with other evidence about workers' labor supply behavior as they approach retirement. It is now known that many mature workers, prior to completely withdrawing from the labor force, leave career jobs, which they have held for many years to move into what Ruhm (1990) called "bridge" jobs—positions in which much lower hours are worked relative to the "career" job. This suggests that workers cannot perfectly adjust hours at their career jobs—something which the evidence in this study confirms. In addition, the study shows that because the transition costs associated with switching to new jobs may be high, constraints induce many workers to withdraw completely from the labor force. As is discussed later, it is likely that these transition costs rise with age and/or tenure in the career job.

The study also extends the work on the determinants of retirement. Most research by economists on the causes of retirement has emphasized purely on worker-side factors such as the effect of age or poor health. Another strand of the literature examines how the various features of public programs, such as Supplemental Security Income or Social Security Disability Insurance, affect retirement propensities. Relatively little work examines the extent to which job characteristics affect when individuals leave the labor force. The results presented below show that at least one of these job features may be a key determinant of retirement.

The next section presents a simple two-period model of labor supply and retirement under hours constraints. Section III discusses the empirical implementation. Our data are discussed and summary statistics presented in Section IV. Various empirical estimates and robustness tests are presented in Section V. Section VI concludes.

II. SIMPLE THEORETICAL FRAMEWORK

Consider workers who in any period t derive utility from leisure and income and from an idiosyncratic taste shifter, ε_i. Let workers be endowed with a positive amount of time and let them consume some of it as leisure. Their income is their wage rate multiplied by the remainder, their hours of work.[3] Since both leisure and income are functions of hours worked, we express utility as a function of hours worked in period t, h_{it}, and the taste shifter, ε_i, so that the utility function is of the form, $U^i(h_{it}, \varepsilon_{it})$. We assume that utility is strictly concave in hours of work.

For any value of the various parameters that determine labor supply, there is an hours level for which the person's utility is higher than would be true for any alternative level of labor supply. We denote this hours level by $h^*_{it} \geq 0$ and call it the person's *desired labor supply* in period t. By definition, this desired hours level is given implicitly by the condition $U^i_h(h^*_{it}, \varepsilon_{it}) = 0$, where U^i_h is the partial derivative of utility, U^i, with respect to hours. Taking a second-order Taylor series expansion around the point h^*_{it}, the utility that an individual receives when he works any arbitrary hours level, h_{it}, is thus:

$$(1)\quad U^i(h_{it},\varepsilon_i) = U^i(h^*_{it},\varepsilon_i) + (h^*_{it} - h_{it})U^i_h(h^*_{it},\varepsilon_i) + (h^*_{it} - h_{it})^2 U^i_{hh}(h^*_{it},\varepsilon_i)$$

Since the second right-hand side term of expression (1) equals 0 by the definition of the desired hours level, h^*_{it}, and since U^i_{hh}, the second derivative of utility with respect to hours is strictly negative because of the assumed concavity of the utility function; the expression simplifies to:

$$(2)\quad U^i(h_{it},\varepsilon_i) = U^i(h^*_{it},\varepsilon_i) - \kappa(h^*_{it} - h_{it})^2$$

where κ is some positive constant. In other words, the utility that individuals derive from working any particular amount of hours equals the utility they would receive from working their desired hours *minus* an amount that is increasing in the extent to which their actual hours deviate from their desired hours. Thus, the quantity, $\kappa(h^*_{it} - h_{it})^2$ can be thought of as the disutility that an individual gets when he is overemployed (underemployed) and working more (less) hours than desired.

Consider now an economy in which workers, with preferences of the sort described above, work for two periods. All firms in the economy pay a worker with observed characteristics, X_i, an hourly wage of $w(X_i)$. Firms differ with respect to the minimum number of hours they require their employees to work each week in order to receive the market wage. Some firms have no minimum hours requirement at the relevant wage, so workers can, from one time period to the next, change their hours to any level they wish. Other firms require their employees to work at least a specific level of hours in each period in order to receive the market wage. At these hours-constrained firms, usual weekly hours cannot be lowered over time to a level below the required amount, so a worker must change jobs if he wishes to work less than the required amount.

At the start of the first period, suppose that workers know how many hours they wish to work in Period 1. Assume that workers take a job that offers them that particular hours level and lets them be initially ignorant of whether the firm is hours constrained or not. Instead, suppose that workers learn the firm's type sometime during the first period of work. Thus, a firm's hours-constrained status is an experience good—something to be learned only after the worker has been employed for at least one period. We focus on workers' decisions between Periods 1 and 2. Each worker must choose whether to (a) remain at their incumbent (first period) employer, (b) switch to a new firm after paying a transition cost of θ, or (c) withdraw completely from the labor force (i.e., retire)

3. We ignored nonlabor income in this framework but incorporated it into our empirical analysis.

between the first and the second periods. In the remainder of the analysis, we focus on workers who work h'_{i1} hours in the first period at their incumbent employers and whose desired hours of work in Period 2 are $h^*_{i2} \geq 0$.

Consider first the case where the worker is *not* hours constrained at the incumbent employer. In this case, the worker's Period 2 labor supply is immediately obvious. If his desired Period 2 hours, h^*_{i2}, equals 0, then the worker retires. If he wishes to work some positive amount in Period 2 different from h'_{i1} hours, then he simply adjusts his labor supply to the new desired level at the incumbent firm since there are no restrictions on any such adjustments at unconstrained jobs. For workers whose hours are flexible, there is thus no difference between the desired and the actual change in hours over time. The worker retires if and only if working 0 h in the later period brings him greater utility than working any positive hours level—that is, if his desired hours in the later period is 0.

Now consider a worker whose minimum hours at the Period 1 job are not flexible downward. If $h^*_{i2} = 0$, then the worker prefers retirement over any option that involves positive labor supply and, as in the unconstrained case, will choose to retire. If $h^*_{i2} > h'_{i1}$, the worker will remain at the first period employer and increase his hours of work. If $0 < h^*_{i2} < h'_{i1}$, however, the worker has options to consider. The worker can remain at the incumbent firm during Period 2 and work h'_{i1} hours. By expression (2), his second period payoff would be:

$$U^i(h^*_{i2}, \varepsilon_i) - k(h^*_{i2} - h'_{i1})^2 \tag{3}$$

if he were to choose this option. A second option is to move to a new employer, at which he could work h^*_{i2}. Since a transition cost of θ would be incurred in the event of such a move, his Period 2 payoff from this second option would be:

$$U^i(h^*_{i2}, \varepsilon_i) - \theta \tag{4}$$

Finally, he could retire, in which case, his Period 2 payoff will be:

$$U^i(h^*_{i2}, \varepsilon_i) - \kappa(h^*_{i2})^2 \tag{5}$$

Clearly, the option selected by the hours-constrained worker depends on which of the three quantities $\kappa(h^*_{i2} - h'_{i1})^2$, θ, or $k(h^*_{i2})^2$ is smallest. Expressions (3–5) summarize the study's main result: Despite the fact that the worker desires to work some *positive* amount, though smaller than in the first period, he would nonetheless *retire* in the second period if:

$$h^*_{i2} \leq \min(h'_{i1}/2, \sqrt{\theta/\kappa}) \tag{6}$$

We assume henceforth that $\min(h'_{i1}/2, \sqrt{\theta/\kappa}) = \sqrt{\theta/\kappa}$ and that $\kappa = 1$.[4]

The forgoing implies that in a sample of individuals working a given amount of hours in Period 1, the probability that a worker whose hours are flexible at the incumbent employer will retire by Period 2 is:

$$\Pr[h^*_{i2} = 0] \tag{7}$$

By contrast, among workers whose usual weekly hours cannot be lowered, the probability of retirement by Period 2 is:

$$\Pr[h^*_{i2} \leq \theta] = \Pr[h^*_{i2} = 0] + \Pr[h^*_{i2} > 0 \cap h^*_{i2} \leq \theta] \tag{8}$$

So long as the second term of the right-hand side expression is non-zero, the probability of retirement from one period to the next should be greater among workers who were not free to vary their hours in the earlier time period.

III. EMPIRICAL IMPLEMENTATION

To empirically implement the model mentioned in the previous section, we suppose that an individual's desired hours of work in any period, h^*_{it}, may be written as:

$$h^*_{it} = \begin{cases} w(X_i) - \beta X_{it} - \delta h_{i,-t} - \varepsilon_i, & \text{if } w(X_{it}) - \beta X_{it} - \delta h_{i,-t} > 0, \\ 0, & \text{otherwise.} \end{cases} \tag{9}$$

4. Both assumptions are for ease of exposition. Our main results are not affected by them in any way.

In expression (9), X_{it} is a vector of the person's observed characteristics, such as age, gender, and educational attainment. The person's wage, $w(X_{it})$, is a function of observable traits. The term $h_{i,-t}$ refers to the person's hours of work in the period previous to t. As mentioned before, ε_i is a person-specific taste shifter, which we assume henceforth is drawn from the standard normal distribution. Expression (9) implies that the quantity $\beta X_{it} + \delta h_{i,-t} + \varepsilon_i$ is the person's reservation wage and is easily generated from many standard utility functions assumed in the labor supply literature.

Let the binary variable $C_{i,-t}$ equal 1 if in the time period previous to t the person works at a job in which it is not possible to alter hours of work, and let the binary variable R_{it} equal 1 if a person employed in the previous time period retires by time period t. From expressions (7) and (8), the probability that R_{it} equals 1 can be written as:

$$(10) \qquad \Pr[R_{it} = 1] = \Pr[h^*_{it} \leq C_{i,-t} * \sqrt{\theta}]$$

Using expression (9), this probability can be written as:

$$(11) \quad \Pr[R_{it} = 1] = \Pr[w(X_{it}) - \beta X_{it} - \delta h_{i,-t} - C_{i,-t} * \sqrt{\theta} \leq \varepsilon_i]$$

Because ε_i is drawn from a standard normal distribution, expression (11) can be written as:

$$(12) \quad \Pr[R_{it} = 1] = 1 - \Phi(w(X_{it}) - \beta X_{it} - \delta h_{i,-t} - C_{i,-t} * \sqrt{\theta}),$$

which implies that:

$$(13) \quad \Pr[R_{it} = 1] = \Phi(w(X_{it}) - \beta X_{it} - \delta h_{i,-t} - C_{i,-t} * \sqrt{\theta})$$

where Φ is the cumulative distribution function of the standard normal distribution.

These two expressions are the contribution of the retired and nonretired individuals to the probit likelihood function. So, to test the model's prediction that working in an hours-constrained job in one period raises the probability of retirement in a later period, we estimate the following probit system by maximum likelihood:

$$(14) \qquad R^*_{it} = \alpha_1 Z_{it} + \alpha_2 C_{i,-t} + \mu_{it}$$

$$R_{it} = 1 \quad \text{if} R^*_{it} > 0.$$

where Z_{it} is a vector of observable controls, including the wage rate, hours of work in the previous period, age, health, and other individual and job-specific characteristics; $C_{i,-t}$ indicates whether the person is unable to vary hours at the job held in the previous period. The coefficient α_2 should be strictly positive if the model is correct.

IV. DATA, SUMMARY STATISTICS, AND BASIC FACTS

To conduct a test of the proposition outlined above, a data source that provides information on workers who are on the cusp of the retirement decision is required. Apart from standard demographic information and information on the usual features of the worker's jobs, there obviously should be information about whether the worker is free to adjust his usual hours of work. Finally, since the model described above concerns a dynamic dimension of labor supply, the data source should be longitudinal in nature. The HRS meets all these requirements.

The HRS is a nationally representative longitudinal survey of households in which there is at least one person born between 1931 and 1941. The first wave of the survey was conducted in 1992. After the first wave, the same sample is surveyed approximately every 1 yr. The age range of the sample over the first three waves of the survey ensures that the sample is at the ideal age to study the determinants of retirement.

We used data from the first three waves of the survey. The sample we analyzed was restricted to persons (a) aged 55–64 yr in Wave 1, (b) who work for pay in Wave 1 but are not self-employed, and (c) who respond to the survey in both Waves 2 and 3. Observations were dropped from the sample if information on age, race, educational attainment, or health status was missing. We also required that there be basic job information on the individual's Wave 1 employment, such as wage rate, usual hours of work, occupation, and union status. If this information was missing, then the respondent was dropped from the sample. Finally, if information on the individual's hours-constrained status was missing, then the individual was dropped from the sample. We discuss this key variable in greater detail below.

Importantly, the analysis sample is restricted to persons meeting the conditions mentioned above who worked at least 20 h/wk in the initial wave. We impose this restriction partly because we wish to study how hours constraints affect the retirement behavior of individuals with reasonably strong initial labor force attachment. For example, we wished to drop from our sample the individuals who are in the midst of adjusting their labor supply toward retirement when they are initially observed. In additional specifications, we strengthen this requirement. In particular, we restricted our sample to individuals who work 30 or more, 35 or more, and 40 or more hours per week in Wave 1. As is seen later, the estimates from samples with these alternative restrictions generally find a stronger relationship between being hours constrained and subsequent retirement, the stronger the person's initial attachment to the labor force.

Table 1 presents means and standard deviations of variables included in our analysis sample. Both the men and the women in the sample were, on average, approximately 58 yr old in Wave 1. Only about one-quarter of both men and women rate their health as "excellent," while about 14% describe themselves as being in "fair" or "poor" health.[5] The health figures are likely a result of the age distribution of the sample. The one noticeable difference in a demographic variable between men and women is the difference in the share who report being married with a spouse present. More than 85% of the male sample but just more than 60% of women respond affirmatively. This difference is generated by the age of the sample. Women tend to marry older men and tend to live longer than men, on average. This implies that in a relatively older sample such as this one, women are much more likely than men to be unmarried due to widowing.

The next three variables measure job-related characteristics. Men in the sample earn higher hourly wages than women, on average, and are more likely to belong to a union. Both men and women have been employed by the firms at which they are initially observed for more than a dozen years. For men, average job tenure as of the first wave is nearly 21 yr, while it is more than 15 yr for women. These figures indicate that, on average, individuals in our sample had not begun making labor supply adjustments when they are initially observed in the first wave. Had this been true, then average tenure would be much lower since workers would disproportionately have only recently moved into their current job. It is thus unlikely that workers in this sample have systematically sorted themselves into jobs with and without hours constraints based on unobserved retirement propensities. As is seen later, we estimated models that include only individuals with 15 or more years of tenure on their current jobs. Estimates from these

TABLE 1
Sample Means and Standard Deviations of Selected Variables

Characteristic in 1992	Women (N = 1,144)	Men (N = 1,340)
Age	57.72 (2.06)	58.24 (2.45)
Married or spouse present	0.61 (0.49)	0.87 (0.33)
White	0.78 (0.41)	0.83 (0.37)
Years of schooling		
<12	0.25 (0.43)	0.29 (0.45)
>12	0.36 (0.48)	0.39 (0.49)
Health		
Excellent	0.25 (0.43)	0.24 (0.43)
Very good	0.30 (0.46)	0.32 (0.47)
Good	0.31 (0.46)	0.31 (0.46)
Fair	0.13 (0.33)	0.10 (0.30)
Poor	0.02 (0.14)	0.02 (0.15)
Union member	0.23 (0.42)	0.30 (0.46)
Job tenure (yr)	15.52 (11.64)	20.91 (13.02)
≤5	0.22 (0.41)	0.14 (0.34)
≥15	0.45 (0.50)	0.63 (0.48)
Log of hourly wage	2.23 (0.58)	2.59 (0.60)
Any nonlabor income	0.45 (0.50)	0.47 (0.50)

Notes: Data are from the first wave of the HRS. Within a small number of cases, they correspond to our analysis samples. See text for additional details.

5. Health information is taken from a question that asks respondents to rate their own health. Legitimate responses include excellent, very good, good, fair, and poor.

models suggest that this type of sorting is not responsible for our main estimates.

With respect to information on hours constraints, respondents were asked the following in the first wave of HRS:

> Not counting overtime hours, could you reduce the number of paid hours in your regular work schedule?

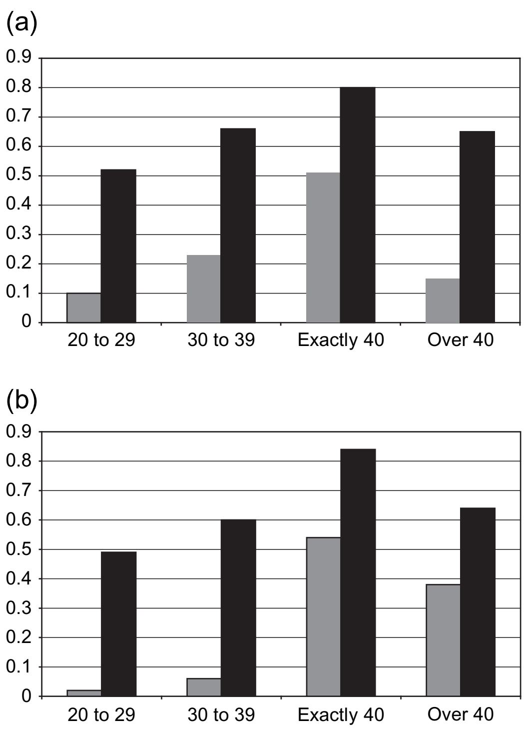
Responses to this question are the key variable in our analysis. A negative response to this question is coded as meaning that the worker is hours constrained and thus unable to lower his regular hours of work at his Wave 1 employer.

Figures 1a,b summarize workers' Wave 1 responses to this question by gender. The figures split workers into four usual weekly hours categories as of the initial wave: those usually working between 20 and 29 h/wk, those working between 30 and 39 h/wk, those working exactly 40 h/wk, and those working more than 40 h/wk. The figures show the fraction of all workers who fall into a particular hours category and, within a given category, the fraction who cannot lower their usual hours of work.

Consider first the distribution of usual weekly hours of work. The figures illustrate the familiar fact that actual hours of work are quite lumpy, with many people located at exactly 40 h/wk. In Figure 1a, for example, just about half of working women in the sample usually work exactly this amount each week. For men, shown in Figure 1b, the fraction working exactly 40 h/wk is closer to 60%. As noted earlier, we studied persons working at least 20 h/wk, so it is not too surprising that labor force attachment, while somewhat stronger for men, is strong for both genders in our sample. Sixty-six percent of women work 40 h or more in Wave 1, while the corresponding fraction for men is about 90%. Nevertheless, many women work less than the traditional "full time" amount: Slightly more than 20% of the women work between 30 and 39 h and about 10% work between 20 and 29 h, while only 8% of men work 39 or fewer hours per week.

The figures also show that across all hours levels, employer-imposed hours constraints, which prevent workers from lowering their usual hours of work, are prevalent for both men and women and tend to increase in prevalence at greater levels of usual weekly labor supply. For example, while 50% of women working between 20 and 29 h/wk cannot lower their usual hours of work, this fraction hours constrained grows to about 80% for those women working exactly 40 h/wk. Sixty percent of women in the other 2 h worked categories are similarly constrained. For men, 50% of those in the lowest hours category cannot adjust hours, and nearly 85% those men working exactly 40 h are similarly constrained. The figures show the ubiquity of hours constraints that prevent reductions in usual weekly hours worked in the U.S. labor market. In the next section, we explore the impact of these constraints on subsequent retirement behavior.

FIGURE 1
Distribution of Weekly Work Hours and Prevalence of Hours Constraints in (a) women and (b) men



Notes: Fraction "Weekly Hours Worked" to left of fraction "Hours-Constrained."

V. RESULTS

A. Simple Difference Estimates

We begin the analysis of the effects of hours constraints on subsequent retirement behavior with some simple difference estimates. The outcome variable here and in all the work that follows is a binary variable that denotes whether the individual is "retired" as of the third wave of the HRS, conducted in 1996. Assessing whether workers are permanently and completely withdrawn from the labor force is difficult. Survey respondents may equate being retired with the receipt of benefits or with having left a job held for much of their working life. Thus, people who are labor force participants might classify themselves as retired, while other individuals whose labor force attachment has ended may not use this expression to describe themselves. Because of such complications, we label workers as retired in this study when they (a) are not working for pay as of the survey date, (b) are not actively seeking work as of the survey date, and (c) have not worked for 1 yr prior to the survey date. Throughout, we will be interested in whether workers are retired by our definition as of the third wave of the survey.

Tables 2 and 3 report 1996 (Wave 3) retirement rates by whether the worker was hours constrained in the job held in 1992 (Wave 1). In each table, the first row pertains to the entire sex-specific sample and subsequent rows show relevant means by the person's age in 1992. Reading down the first two columns shows that, irrespective of hours constraints, retirement rates are generally increasing with age. We have grouped those aged 62 yr and older due to small cell sizes. The last column shows the difference in retirement rates across those who are and are not hours constrained, by particular ages and for the entire sex-specific sample.

Table 2 shows that overall, women who faced hours constraints were 6.4% points more likely to be retired 4 yr later than those who were not and this difference is statistically significant. Given a mean retirement rate of about 24%, the difference in retirement rates suggests roughly a 25% increase in the probability of retirement due entirely to being constrained. The age-specific numbers show that this effect is generated primarily from women aged less than 59 yr in the initial wave. This is not surprising as the social security system has large discontinuous retirement incentives that make labor force withdrawal very likely at ages 62 and older. These conditional means show that on average, facing hours constraints raises the likelihood women will withdraw from the labor force before the traditional early retirement age.

Table 3 presents the corresponding simple difference estimates for men. As with women, retirement rates rise with age, regardless of constraint status. The first row of the table shows that previously hours-constrained men have a systematically higher probability of retiring. The statistically significant 7.3% point difference in retirement probabilities translates into about a 25% increase in retirement probability. Like those for women in our sample, some of the age-specific results are of the expected sign but not statistically different from 0. An interesting difference between

TABLE 2
Fraction of Working Women in 1992 Retired by 1996, by Ability to Lower Usual Hours in 1992

	Fraction Retired in 1996		
	Free to Lower Hours in 1992	Cannot Lower Hours in 1992	Difference in Fraction Retired
Overall	0.195 (0.022)	0.259 (0.015)	−0.064 (0.027)
Age in 1992 (yr)			
55	0.019 (0.019)	0.109 (0.026)	−0.090 (0.032)
56	0.092 (0.036)	0.136 (0.031)	−0.044 (0.048)
57	0.063 (0.035)	0.138 (0.030)	−0.075 (0.046)
58	0.174 (0.057)	0.285 (0.041)	−0.111 (0.070)
59	0.371 (0.083)	0.366 (0.048)	0.005 (0.096)
60	0.371 (0.083)	0.407 (0.053)	−0.036 (0.098)
61	0.500 (0.100)	0.506 (0.056)	−0.006 (0.115)
62 or older	0.556 (0.176)	0.522 (0.112)	0.034 (0.202)

Notes: Individuals aged 62 yr and older were grouped together due to relatively small cell sizes. Data are from multiple waves of the HRS. Standard errors are given in parentheses. See text for additional details.

TABLE 3
Fraction of Working Men in 1992 Retired by 1996, by Ability to Lower Usual Hours in 1992

	Fraction Retired in 1996		
	Free to Lower Hours in 1992	Cannot Lower Hours in 1992	Difference in Fraction Retired
Overall	0.274 (0.025)	0.347 (0.015)	−0.073 (0.029)
Age in 1992 (yr)			
55	0.077 (0.037)	0.123 (0.027)	−0.046 (0.046)
56	0.071 (0.040)	0.166 (0.030)	−0.095 (0.050)
57	0.156 (0.055)	0.220 (0.035)	−0.064 (0.065)
58	0.184 (0.064)	0.264 (0.039)	−0.080 (0.075)
59	0.333 (0.088)	0.422 (0.046)	−0.089 (0.099)
60	0.325 (0.075)	0.573 (0.047)	−0.248 (0.089)
61	0.438 (0.089)	0.656 (0.049)	−0.218 (0.102)
62 or older	0.653 (0.069)	0.656 (0.050)	−0.003 (0.085)

Notes: Individuals aged 62 yr and older were grouped together due to relatively small cell sizes. Data are from multiple waves of the HRS. Standard errors are given in parentheses. See text for additional details.

women and men is where in the age profile significant differences are found. In particular, the largest differences by constraint status are found at ages 60 and 61.

Despite the fact that some of the age-specific results are not statistically different from 0, the overall sex-specific differences are each practically large. Moreover, the age-specific differences are largely consistent with the theory presented earlier. The next natural question is whether these estimates can be interpreted as being causal. We briefly examined whether observable traits are similar for constrained and unconstrained workers. This comparison gives us two things. First, the extent to which observables differ between the groups is a good indicator of the gains from a more formal framework such as that presented in the next subsection. Second, if observable factors are the same between hours-constrained and unconstrained workers, then it is more likely that the results shown above are not generated from differences in unobserved heterogeneity. In effect, similarity in the means for the two groups suggests "random assignment" of hours-constrained status.

Table 4 presents means for demographic and job-related variables by constraint status in the initial wave. The first set of numbers indicates that for women, there are no differences in any of the demographic variables between workers by their hours-constrained status. For men, the only demographic variable for which there is a difference between the two groups is in completed schooling: Hours-constrained men tend to have less formal education. With respect to the job-related variables, the only factor which, for both men and women, differs significantly between those with and without hours constraints is the degree of unionization. Hours-constrained workers are much more likely to be union members.

Because there are differences in education and union status between the two types of workers, it seems unlikely that constraint status is assigned randomly. The simple difference estimates thus cannot be interpreted as causal estimates. In particular, part of the difference between the average retirement behaviors of the two types of workers could stem not from their hours-constrained status but rather from the fact that something about unionized jobs or being more highly educated. In the analysis that follows, we controlled for these effects in models whose results are presented in the next section. The estimates also suggest that there could be latent differences between constrained and unconstrained workers, possibly due to systematic sorting into constraint type. Later, we performed some simple robustness checks to account for this possibility.

B. Regression Results

In this section, we assess the effect of hours constraints on subsequent retirement using the probit model outlined above. The estimated coefficients from the probit do not measure marginal effects of a slight change in the value of the associated right-hand side variable. We therefore present the mean marginal effect for our variable of primary interest, hours-constrained status, in addition to its

TABLE 4
Means and Standard Errors of Key Variables, by Sex and Ability to Lower Usual Hours Worked in 1992

	Women				Men			
	Constrained		Unconstrained		Constrained		Unconstrained	
	Mean	Standard Error	Mean	Standard Error	Mean	Standard Error	Mean	Standard Error
Demographics								
Age	57.78	0.06	57.73	0.10	58.18	0.07	58.55	0.12
White	0.76	0.01	0.80	0.02	0.82	0.01	0.85	0.02
Married	0.62	0.02	0.62	0.02	0.86	0.01	0.87	0.02
Health								
Excellent	0.24	0.01	0.25	0.02	0.23	0.01	0.26	0.02
Very good	0.29	0.01	0.30	0.02	0.31	0.01	0.33	0.02
Good	0.32	0.01	0.31	0.02	0.33	0.01	0.29	0.02
Fair	0.13	0.01	0.12	0.02	0.10	0.01	0.10	0.01
Poor	0.02	0.00	0.02	0.01	0.03	0.00	0.03	0.01
Schooling (yr)								
>12	0.37	0.02	0.34	0.02	0.34	0.01	0.46	0.02
<12	0.25	0.01	0.46	0.02	0.30	0.01	0.24	0.02
Any nonlabor income	0.46	0.02	0.43	0.02	0.46	0.01	0.51	0.02
Job characteristics								
Union member	0.27	0.01	0.17	0.02	0.37	0.01	0.17	0.02
Log of hourly wage	2.27	0.02	2.10	0.03	2.58	0.02	2.61	0.03
Job tenure (yr)	15.84	0.35	15.12	0.63	21.03	0.36	20.90	0.66
≤5	0.19	0.01	0.27	0.02	0.13	0.01	0.17	0.02
≥15	0.46	0.02	0.40	0.02	0.63	0.01	0.61	0.02

Notes: Constrained individuals reported that they cannot lower their usual weekly hours of work, while unconstrained individuals reported that they are able to do so. Data are from the first wave of the HRS. See text for additional details.

probit coefficient and associated standard error. The mean marginal effect we present is the average of individual marginal effects rather than the marginal effect evaluate at the means of the exogenous variables.[6]

Table 5 presents the probit results for women. For each variable, the table presents the probit coefficient and its standard error in parentheses. The mean marginal effect, described directly above, is also included in brackets for the hours-constrained variable. The first column of Table 5 is a base specification, in which retirement in 1996 is related to standard demographic controls and to controls for hours of work in the initial wave. We controlled for previous hours using a series of dummy variables indicating whether the person worked 20–29, 30–39, exactly 40, or more than 40 h/wk.

Whereas differences in race and education produce no statistically different effect on retirement probabilities, women who are married with a spouse present are more likely to retire than their single counterparts. Likewise, those who reported in 1992 that they were in fair or poor health had a retirement probability that is about 8% points higher than those who reported excellent health.[7]

The regression displayed in the second column adds to this base specification the variable of greatest interest—a dummy variable measuring whether each worker was unable to lower his usual hours of work at the job he held as of the initial wave. Throughout, we continue to refer to such individuals as being hours constrained. The estimates indicate that

6. This is the estimate recommended by Chamberlain (1982). The mean marginal effect is the mean across the sample of the quantity $\beta_j\phi(\beta X_i)$, where β_j is the estimated probit coefficient for the characteristic x_j, $\phi(.)$ is the marginal density of the Standard Normal distribution, and X_i measures the full vector of characteristics of person i.

7. Individuals in fair or poor health represent the omitted group for the set of health status dummies included in the model.

TABLE 5
Probit Estimates of Hours Constraints on Job Held in 1992 on Retirement Probability of Women in 1996

Characteristic in 1992	(I)	(II)	(III)	(IV)
Cannot lower hours at job	—	0.232 (0.098) [0.064]	0.181 (0.102) [0.049]	0.181 (0.102) [0.049]
Married	0.226 (0.090)	0.226 (0.091)	0.190 (0.094)	0.179 (0.095)
White	−0.051 (0.108)	−0.046 (0.109)	−0.062 (0.117)	−0.078 (0.119)
Years of schooling				
<12	−0.047 (0.116)	−0.055 (0.117)	0.084 (0.127)	0.090 (0.128)
>12	0.126 (0.099)	0.123 (0.100)	−0.105 (0.115)	−0.118 (0.116)
Age	0.241 (0.022)	0.239 (0.022)	0.255 (0.022)	0.256 (0.023)
Health				
Excellent	−0.291 (0.150)	−0.297 (0.150)	−0.313 (0.153)	−0.321 (0.153)
Very good	−0.218 (0.141)	−0.234 (0.142)	−0.221 (0.144)	−0.227 (0.145)
Good	−0.148 (0.137)	−0.157 (0.138)	−0.138 (0.139)	−0.141 (0.139)
Union member	—	—	0.167 (0.105)	0.169 (0.105)
Log of hourly wage	—	—	0.154 (0.094)	0.148 (0.094)
Any nonlabor income	—	—	—	0.082 (0.097)
Controls for Wave 1 hours worked?	Yes	Yes	Yes	Yes
Industry and occupation dummies?	No	No	Yes	Yes
Include nonlabor income?	No	No	No	Yes
N	1,144	1,134	1,129	1,129
Dependent variable mean	0.24	0.24	0.24	0.24

Notes: Data are from multiple waves of the HRS. Robust standard errors are given in parentheses and mean marginal effects of hours-constrained status on retirement in square brackets. See text for additional details.

being hours constrained raises the probability of retirement by more than 6% points. This is almost exactly the same point estimate as found in the simple difference estimates from the first row of Table 2. In practical terms, the marginal effect is quite large as it is almost 80% of estimated marginal effect of being in fair or poor health—the factor most commonly thought to be a determinant of early labor force withdrawal.

In the third column of the table, we add job-related variables to determine if the effect we attribute to hours constraints is due instead to other job features that might be correlated with hours flexibility. For example, there may be "good" and "bad," which differ not only with respect to whether persons who work there are free to lower their hours of work but also with respect to other unobserved dimensions like forms of compensation of work conditions. If hours-constrained jobs are bad jobs, which workers are eager to leave for *other* reasons, we will incorrectly attribute the entire increased retirement associated with such jobs to the constraints. In principle, job quality is a multidimensional phenomenon, which is difficult to measure, but we argue that the relevant dimensions of quality are likely well summarized by a few key variables. In particular, the model presented in Column III controls for the log of hourly wage, union status, and a complete set of two-digit industry and occupation codes.[8] Estimates from this model suggest that higher wages and being in a union both raise the probability of subsequent retirement for women. More to the point, adding these job characteristics lowers the estimated impact of the hours-constrained variable by about one-quarter, though the impact remains marginally statistically significant for women.[9]

8. Because information on some of these variables is missing for a few sample members, sample size falls slightly moving from Column II to Column III in Tables 5 and 6 (women and men, respectively).

9. Though not reported separately, the log wage variable is responsible for most of this reduction in the implied effect of hours-constrained status.

Column IV adds a measure of nonlabor income to the previous column's model.[10] In particular, we included an indicator that equals 1 if the individual had *any* nonlabor income in the first wave. We used this discrete measure of nonlabor income since less than half of the respondents in our sample reported receiving such income. We found little evidence that the receipt of nonlabor income influences the retirement decisions of women. More relevant to our hypothesis, we found that its inclusion does not impact the estimated effect of being hours constrained on retirement. Indeed, moving from Column III to Column IV in Table 5, the relevant estimates are identical.

In Appendix 1, we report estimates from models that include nonlabor income as a continuous variable. We also included various measures of wealth on the assumption that wealth may also influence retirement decisions even if it does not directly generate disposable income. As can be seen in the first column of Appendix 1, estimates of the impact of being hours constrained differ very little from the those presented in Table 5. In particular, the estimated effect of hours-constrained status on retirement ranges from 4.6 to 5.1% points.

Table 6 presents probit estimates for men. Overall, the base results presented in the first column are similar to those for women. Race and education levels have no significant impact on retirement probabilities. Again, health status seems to play an important role in determining subsequent retirement as individuals who report fair or poor status are between 6% and 7% more likely to retire than their counterparts who report excellent health. Interestingly, the base specification suggests that, unlike women, there is no systematic relationship between marital status and retirement. A likely explanation for this gender difference is the age gap between typical spouses and the fact that many spouses wish to retire jointly. For example, spouses of women sampled are much more likely to be retired than spouses of sample men, on average.

In Column II, the hours-constrained variable is once again added to the specification. As with women, the inability to lower the hours of work is estimated to have a substantial impact on retirement. Indeed, we found that being hours constrained raises one's probability of retirement by slightly more than 10% points in this specification. This exceeds the simple difference estimates and is among the largest estimated marginal effects we estimated. As seen in Column III, when we add controls for the hourly wage, union status, and a set of industry and occupation dummies, the estimated effect of being hours constrained is about two-thirds as large—still a practically large effect that remains precisely estimated.[11] As with women, we add a binary variable for the receipt of nonlabor income in Column IV.

Contrary to our estimates for women, we found evidence that nonlabor income receipt increases the probability of retirement among men. Again, we found that the inclusion of the nonlabor income dummy does not impact the estimated impact of hours-constrained status. In fact, it raises the estimated impact very slightly. We again estimated the model in Column IV, with different measures of nonlabor income and wealth, and as can be seen in Appendix 1, the estimates remain quite similar, as was the case for women. In particular, the estimated effect of hours-constrained status on retirement for men ranged from 6.6% to 7.2% points.

C. Additional Regression Results

As noted earlier, the analysis presented thus far restricts attention to individuals working at least 20 h/wk as of the first wave. We imposed this restriction to reduce the chance that the sample included individuals who had already begun the retirement process when first observed in the data. If persons in the midst of retirement are included in the analysis, and if hours constraints are more binding at lower levels of weekly hours, the foregoing results will *overstate* the impact of being hours constrained.[12] To address this concern, we impose more restrictive restrictions

10. Nonlabor income includes the following income types: household business income, farm income, business income, gross rent, dividend and interest income, trust funds, royalties, and other asset income. Information on nonlabor income and various measures of wealth, discussed below, are taken from cleaned files provided to the HRS by the Rand Corporation.

11. Though not reported separately, union status is responsible for most of this reduction in the implied effect of hours-constrained status.

12. We thank an anonymous referee for pointing this out to us.

TABLE 6
Probit Estimates of Hours Constraints on Job Held in 1992 on Retirement Probability of Men in 1996

Characteristic in 1992	(I)	(II)	(III)	(IV)
Cannot lower hours at job	—	0.315 (0.094) [0.105]	0.200 (0.098) [0.067]	0.207 (0.098) [0.069]
Married	−0.078 (0.118)	−0.101 (0.119)	−0.109 (0.124)	−0.132 (0.124)
White	0.135 (0.103)	0.129 (0.104)	0.146 (0.107)	0.119 (0.108)
Years of schooling				
<12	−0.091 (0.098)	−0.107 (0.099)	0.026 (0.104)	0.047 (0.105)
>12	0.011 (0.092)	0.024 (0.092)	0.040 (0.104)	0.036 (0.104)
Age	0.230 (0.017)	0.238 (0.017)	0.255 (0.017)	0.257 (0.017)
Health				
Excellent	−0.189 (0.135)	−0.143 (0.138)	−0.227 (0.144)	−0.235 (0.145)
Very good	−0.065 (0.127)	−0.040 (0.129)	−0.106 (0.131)	−0.114 (0.133)
Good	−0.039 (0.125)	−0.025 (0.127)	−0.072 (0.130)	−0.072 (0.131)
Union member	—	—	0.417 (0.090)	0.416 (0.090)
Log of hourly wage	—	—	0.379 (0.095)	0.342 (0.096)
Any nonlabor income	—	—	—	0.206 (0.087)
Controls for Wave 1 hours worked?	Yes	Yes	Yes	Yes
Industry and occupation dummies?	No	No	Yes	Yes
Include nonlabor income?	No	No	No	Yes
N	1,340	1,324	1,317	1,317
Dependent variable mean	0.33	0.33	0.33	0.33

Notes: Data are from multiple waves of the HRS. Robust standard errors are given in parentheses and mean marginal effects of hours-constrained status on retirement in square brackets. See text for additional details.

on initial work hours. Specifically, we reestimate the main models with the sample restricted to persons who in the first wave work (a) 30 or more hours per week, (b) 35 or more hours per week, and (c) 40 or more hours per week. Table 7 displays results from these alternative samples. Models estimated correspond to the fullest specifications in the fourth column of Tables 5 and 6.

The results indicate that the estimated impact of hours constraints on subsequent retirement is larger, the more restrictive the initial work restriction imposed on the sample. In particular, the estimated impact of being hours constrained for men grows from 6.9% points, with the 20 or more hours restriction, to 8.3% points, with the 40 or more hours restriction. Corresponding figures for women are 4.9% and 7.3% points, respectively. These results suggest that the 20-h restriction does not impart an upward bias in the estimated effect of being constrained.

While we are interested in the impact of being hours constrained on retirement, the model in Section II can easily be extended to show that facing hours constraints should raise the likelihood that a worker changes jobs. Table 8 presents estimates from a model restricted to persons working in both the first and the third waves of the HRS. The outcome variable is binary and denotes whether the worker moves to a new employer between the first and the third waves. We found that for men, the estimated effect is positive and marginally statistically significant. For women, the sign of the hours-constrained coefficient is positive but not statistically different from 0.

Overall, the probit estimates confirm the simple difference estimates for both men and women, as relatively large increases in retirement probability are found for being hours constrained, even after detailed individual and job characteristics are held constant. For both men and women, the results suggest that some of the hours-constrained effect is attributable to the fact that certain kinds of jobs, such as unionized positions, tend to feature hours constraints and also tend to be jobs from which workers retire sooner. However,

TABLE 7

Effect of Hours Constraints on Subsequent Retirement, by Hours Worked per Week in 1992

	20+ h/wk		30+ h/wk		35+ h/wk		40+ h/wk	
	Women	Men	Women	Men	Women	Men	Women	Men
Cannot lower hours	0.181 (0.102)	0.207 (0.098)	0.203 (0.109)	0.212 (0.100)	0.263 (0.122)	0.229 (0.103)	0.270 (0.134)	0.252 (0.106)
	[0.049]	[0.069]	[0.056]	[0.071]	[0.070]	[0.076]	[0.073]	[0.083]
Union	0.169 (0.105)	0.416 (0.090)	0.115 (0.110)	0.413 (0.091)	0.138 (0.115)	0.434 (0.092)	0.191 (0.128)	0.458 (0.094)
	[0.050]	[0.148]	[0.034]	[0.147]	[0.040]	[0.154]	[0.057]	[0.163]
Log of wage	0.148 (0.094)	0.342 (0.096)	0.184 (0.104)	0.347 (0.100)	0.164 (0.110)	0.353 (0.103)	0.156 (0.121)	0.351 (0.106)
	[0.042]	[0.110]	[0.053]	[0.119]	[0.046]	[0.121]	[0.045]	[0.120]
Nonlabor income	0.082 (0.097)	0.206 (0.087)	0.079 (0.102)	0.199 (0.088)	0.088 (0.106)	0.188 (0.089)	0.044 (0.116)	0.188 (0.092)
	[0.023]	[0.071]	[0.023]	[0.068]	[0.025]	[0.065]	[0.013]	[0.064]
N	1,129	1,317	993	1,291	909	1,268	738	1,227
Dependent mean	0.24	0.33	0.24	0.33	0.24	0.32	0.24	0.32

Notes: Estimates are derived from models with the most extensive set of controls (see Column IV in Tables 5 and 6). The first two columns of this table are taken from Tables 5 and 6, respectively. Robust standard errors are given in parentheses and mean marginal effects in square brackets.

the estimates indicate an increased likelihood of retirement, attributable purely to the presence of constraints for men and women alike.

D. Robustness Tests and Alternative Specifications

In this section, we reestimate versions of the above-mentioned model to assess whether systematic sorting or measurement error can account for our estimates. We describe these two potential problems and then the associated robustness tests we performed.

In the absence of an experimental design that randomly assigns people to jobs with and without hours constraints, there is no way to be sure that the strong effects found in the simple difference and probit models are not attributable to unobserved differences between constrained and unconstrained workers. In particular, it is possible that workers with strong latent retirement propensities may have sorted into hours-constrained jobs. If so, it would be their unobserved preference for retirement, and not their hours-constrained status, driving our findings. The fact that there are few differences, on average, between the two types of workers in the demographic and job characteristics suggests that differences in unobservables are not driving our findings, but there is no way to be absolutely certain about this.[13]

We conducted a simple test to assess the possible importance of this type of sorting. The logic of the test is straightforward. Workers likely sort into jobs throughout their working lives for a variety of reasons, with the importance of any particular reason varying over the life cycle. In particular, when sorting occurs for reasons having to do with retirement, we would expect it to occur relatively later in the life cycle. Thus, among a group of older workers in 1992, those who sorted into their jobs *for reasons having to do with retirement* should have relatively short job tenures in this base year. Workers with long job tenures may well have sorted into their jobs in nonsystematic ways, but it is unlikely that they did so with an eye to whether they would want to retire early 15 to 20 yr or longer into the

13. As noted, our regressions control for education and union status, the only observable traits with substantial differences across the two worker types. We also estimated models separately by union status, and the estimates are qualitatively similar to those already presented.

TABLE 8
Probit Estimates of the Effect of Hours Constraints in 1992 on Whether Worker Moves to a New Job by 1996

Characteristic in 1992	Women	Men
Cannot lower hours at job	0.098 (0.124) [0.024]	0.208 (0.122) [0.054]
Married	−0.014 (0.117)	−0.213 (0.155)
White	−0.100 (0.139)	0.331 (0.154)
Years of schooling		
<12	−0.306 (0.155)	−0.298 (0.146)
>12	−0.117 (0.143)	−0.152 (0.131)
Age	0.012 (0.028)	0.026 (0.023)
Health		
Excellent	0.038 (0.209)	−0.226 (0.188)
Very good	0.248 (0.197)	−0.012 (0.175)
Good	−0.007 (0.194)	−0.152 (0.178)
Union member	−0.340 (0.147)	−0.120 (0.126)
Log of hourly wage	−0.308 (0.122)	−0.058 (0.104)
Any nonlabor income	−0.150 (0.117)	0.029 (0.114)
Controls for Wave 1 hours worked?	Yes	Yes
Industry and occupation dummies?	Yes	Yes
N	778	825
Dependent variable mean	0.15	0.14

Notes: Data are from multiple waves of the HRS. Sample restricted to individuals not retired by 1996. Robust standard errors are given in parentheses and mean marginal effect of hours-constrained status on retirement in square brackets. See text for additional details.

future. We estimated the impact of being hours constrained on retirement for workers whose job tenure when first observed in 1992 is substantial.

Our second robustness test assesses the extent to which measurement error in the hours-constrained variable may be driving our estimates. Classical measurement error, in which the error is uncorrelated with the unobserved retirement propensity, would have the effect of attenuating any estimated effects of retirement, thereby making the strong positive estimates presented more convincing. However, it has been shown that measurement error is unlikely to be classical when the mismeasured variable is binary, as in this case.[14] More importantly, regression estimates can be biased where there is mean-reverting error. The regressions presented earlier used only hours-constrained status information from the first wave of the HRS. One way of correcting for possible measurement error is to use information from some other year. In this spirit, we estimated an alternative specification that relates hours-constrained status information from both the first and the second waves to retirement status by Wave 3. Because this test examines workers who were employed in each of the first two waves, the sample is different from that used in the regressions mentioned above.

The three panels of Table 9 present estimates from regressions in which we reestimated the probit models displayed in the last columns of Tables 5 and 6 on three different samples. In Panel A, the model is estimated on a sample of workers who, when they were first observed in the first wave, had worked with their employer for at least 5 yr. Panel A displays the probit coefficient on the hours-constrained dummy variable, its robust standard error, and the implied percentage point effect on retirement probability. The estimates indicate larger implied effect for both men and women, relative to what we presented in Column IV of Tables 5 and 6. In particular, the estimates in Panel A imply that being hours constrained raises the female probability of retirement by 5.6% points and the male retirement probability by 8.7% points.

Panel B of Table 9 restricts our sample to persons who had been with their Wave 1

14. See Kane, Rouse, and Staiger (1999).

TABLE 9
Probit Estimates of the Effect of Hours Constraints on Whether Retired in 1996 for Alternative Samples

	Women	Men
A. Persons with 5+ yr of tenure in 1992		
Dummy: "Cannot reduce hours at job"	0.192 (0.118) [0.056]	0.254 (0.106) [0.087]
Dependent mean	0.26	0.35
N	876	1137
B. Persons with 15+ yr of tenure in 1992		
Dummy: "Cannot reduce hours at job"	0.398 (0.155) [0.121]	0.367 (0.120) [0.130]
Dependent mean	0.30	0.35
N	536	852
C. Persons employed in 1992 and 1994		
Dummy: "Cannot reduce hours in both 1992 and 1994"	0.266 (0.233) [0.063]	0.509 (0.180) [0.130]
Dummy: "Cannot reduce hours in 1992 only"	−0.213 (0.315) [−0.048]	0.347 (0.207) [0.103]
Dummy: "Cannot reduce hours in 1994 only"	−0.343 (0.311) [−0.074]	0.248 (0.224) [0.072]
Dependent mean	0.16	0.23
N	894	984

Notes: Data are from multiple waves of the HRS. Robust standard errors are given in parentheses and mean marginal effects in square brackets. Regression controls are identical to those in the Column IV of Tables 5 and 6. See text for additional details.

employer for at least 15 yr when initially observed. This is the sample for which the likelihood of sorting on latent retirement propensities should be very low. As can be seen, the estimated impact of being hours constrained increases substantially for both women and men, and each effect is based on a very precise coefficient estimate. While neither of these tests completely eliminates the possibility of sorting as an explanation for our main estimates, the strong results in these samples, where sorting seems less likely, suggests a causal role for hours constraints.

Panel C presents estimates from a sample restricted to persons employed in both the first and the second waves. The regression model contains the same controls as models in the last column of Tables 5 and 6. However, rather than the single hours-constrained report, the regression contains three measures: whether the worker reports being hours constrained in both 1992 and 1994, constrained

APPENDIX 1
Estimates of the Effect of Hours Constraints on Subsequent Retirement Controlling for Alternative Measures of Nonlabor Income and Wealth

	Women	Men
Any nonlabor income	0.181 (0.102) [0.049]	0.207 (0.098) [0.069]
Nonlabor income	0.182 (0.103) [0.050]	0.204 (0.098) [0.068]
Any wealth	0.180 (0.102) [0.049]	0.203 (0.098) [0.068]
Total wealth	0.188 (0.103) [0.051]	0.199 (0.099) [0.067]
Total wealth, excluding housing	0.186 (0.103) [0.051]	0.196 (0.099) [0.066]
Change in total wealth	0.171 (0.104) [0.047]	0.216 (0.100) [0.072]
Change in nonhousing financial wealth	0.166 (0.103) [0.046]	0.209 (0.100) [0.070]
Change in nonhousing assets	0.161 (0.103) [0.046]	0.212 (0.100) [0.071]

Notes: The table presents the relevant impact of being hours constrained in 1992 on subsequent retirement in 1996. Information from the first row is taken from Tables 5 and 6, respectively. Change variables are measured from 1992 to 1994. Continuous income and wealth measures are in 1992 dollars. Sample size for first five rows for women is 1,129 and for last three rows is 1,091. Corresponding figures for men are 1,317 and 1,268. Data are from multiple waves of the HRS. Robust standard errors are given in parentheses and mean marginal effects in square brackets. See text for additional details.

only in 1992, and constrained only in 1994. The excluded category is being hours constrained in neither period. For men, the results suggest strongly that measurement error is not driving our main estimates. We found that relative to being unconstrained in both periods, men who report being constrained in both periods are 16% points more likely to retire. The other two categories, which constitute about 10% of persons who ever report being hours constrained, have smaller estimated effects. We found evidence that women who are hours constrained in both periods are about 6% points more likely to retire than their counterparts who are unconstrained in both periods, though this effect is not estimated as precisely as for men. The results suggest that measurement error in hours-constrained status does not account for our main estimates.

VI. CONCLUSION

In this study, we examined how facing hours constraints on the job affects the retirement decisions of mature workers. The subject has received relatively little previous attention in the literature, despite evidence that restrictions on hours of work are a prevalent feature of the labor market and evidence that the presence of these constraints makes it difficult for workers to satisfy changing labor supply preferences without changing jobs.

Using multiple waves of the HRS and using a series of robustness checks and empirical tests, we found that mature hours-constrained workers, and especially men, are much more likely to retire by some future date than are workers who are free to adjust their hours of work. Our main estimates are robust to a series of alternative specifications, suggesting that they are not attributable merely to unobserved job or worker characteristics. Our estimates suggest that the impact of being hours constrained is practically important as the size of this effect approaches the impact of being in poor health on retirement.

It is known that many workers prior to the time they completely withdraw from the labor force "partially retire." That is, they move to jobs at which they work substantially fewer hours. Hours constraints on career jobs may account for these patterns. We found evidence that such constraints also push some mature workers out of the labor force entirely, perhaps because the transition costs of moving to a new job are especially high for older workers. It is tempting to argue that the presence of hours constraints imposes an efficiency loss on the economy, insofar as such rigidities induce individuals who are willing to work some positive amount to instead stop working completely. However, this claim may be invalid as the study, and the larger literature, is silent on the question of *why* hours constraints exist in the first place. Future research should address this question.

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