Labor Supply, Hours Constraints, and Job Mobility

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ABSTRACT

If hours can be freely varied within jobs, the effect on hours of changes in preferences for those who do change jobs should be similar to the effect on hours for those who do not change jobs. Conversely, if employers restrict hours choices, then changes in preferences should affect hours more strongly when the job changes than when it does not change. For a sample of married women we find that changes in many of the labor supply preference variables produce much larger effects on hours when the job changes.

I. Introduction

Most of the literature on labor supply is based on the assumption that workers can freely choose hours at a parametric wage.

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Given this assumption, labor supply has limited implications for mobility from one job to another and for the relationship between job mobility and hours changes. Because hours can be freely varied within jobs, hours do not have an independent effect on job choice once the wage is accounted for. Likewise, the effect of changes in labor supply preferences on hours will not depend on whether a quit occurs.

There are, in fact, strong theoretical arguments and empirical evidence that hours cannot be freely varied within jobs, but are instead strongly influenced by employer preferences.[1] If jobs consist of fixed hours-wage packages, then changes in labor supply preferences will result in hours changes only if the worker changes jobs. Furthermore, if information about job opportunities is imperfect, workers may not always be able to move to jobs with desired hours.[2] In this case, changes in labor supply preferences will result in actual hours changes only if a job offering a superior hours-wage package can be found.

A natural way to test the hypothesis that job mobility is necessary if changes in preferences are to affect hours is to estimate hours-change equations, allowing the effects of changes in indicators of preferences to vary depending on whether or not a quit occurred. If hours can be freely varied within jobs, the effect of changes in preferences on hours for those who do change jobs should be similar to the effect on hours for those who do not change jobs. Conversely, if hours constraints within jobs are important, then changes in preferences should affect hours more strongly when the job changes than when it does not change.

1. See Card (1987) for a survey. On the theoretical side, models of labor demand that include worker specific costs as well as nonlinearities in the relationship between hours per worker and output suggest that work hours may be a job characteristic about which firms have particularly strong preferences. See Lewis (1969) and Deardorff and Stafford (1976). On the empirical side, there is evidence to support the view that the constraints placed by firms on hours choice are quantitatively significant. A number of studies indicate that much unemployment reflects constraints on choice of hours of work. These include Ham (1979, 1982 and 1986) and Moffitt (1984). Gustmann and Steinmeier (1983, 1984) have shown that persons near retirement age often must change jobs to reduce work hours. Kahn and Lang (1987) provide a recent analysis of worker reports of overemployment and underemployment.

2. Rosen (1976), Moffitt (1984), Lundberg (1984), and Biddle and Zarkin (1989) analyze labor supply under the assumption that workers face an hours-wage locus rather than a fixed wage rate, but assume that workers may costlessly locate a job offering the hours-wage combination that they prefer. Abowd and Ashenfelter's (1981) study of compensating differentials for systematic underemployment and for unemployment risk as well as a number of subsequent studies consider the implications of the unemployment risk associated with jobs for their desirability, but abstract from the issue of how workers locate jobs offering the optimal combination of underemployment, unemployment risk and wages given worker's preferences. Siow (1987) estimates a model in which workers may adjust hours within a firm, but face a wage penalty for deviating from the firm's preferred hours level.

We examine this issue by estimating equations of the following form:

$$(1)\quad \Delta^m H_{it} = \Delta^m X_{it}[\alpha + \beta Q_{it}^m] + \varepsilon_{it},$$

where $\Delta^m H_{it}$ is the change in hours (hours/week, weeks/year or hours/year) between time t and $t - m$, $\Delta^m X_{it}$ is the change between t and $t - m$ in a vector of variables measuring labor supply preferences, and Q_{it}^m is a variable which equals 1 if a quit occurred between t and $t - m$, and 0 if there was no quit.

A finding that β is close to 0 and α is nonzero would suggest that the effect of preference changes on hours is independent of whether the job changed, implying hours flexibility within jobs. A finding that α is close to 0 (and β nonzero) supports the hypothesis of hours constraints within jobs. In reality, jobs offer varying degrees of hours flexibility. If hours constraints exist for part of the population, one might expect that α and β will be of the same sign, but that $\alpha + \beta$ will be larger (in absolute value) than α.

We present estimates of Equation (1) for a sample of married women from the Panel Study of Income Dynamics (PSID) who are employed at two points in time. Married women are used for the empirical work because we observe many variables, such as the number and age composition of children and the spouse's work hours and income, which may be important labor supply determinants for this group. They are also of special interest because explanations of the male/female wage gap that emphasize sex differences in the level and variability of work hours implicitly take the view that employers care about work hours. We discuss the data in Section II.

Our basic result in Section III is that changes in many of the labor supply preference variables do have much larger effects on hours when the job changes.[3] Taken at face value, the results suggest hours restrictions are important within jobs. One would expect that if a job consists of an hours wage package, then changes in labor supply preferences may be an important factor in mobility decisions.

We address two alternative interpretations of the basic results theoretically and empirically in Section IV, however, using a simple model of mobility and hours determination. The first alternative is that hours are perfectly flexible within jobs, but labor supply response parameters vary across individuals and happen to be systematically larger for those who

3. We estimated similar equations for men and unmarried women. The results for these groups were disappointing. Basically, few of the variables which might qualify as determinants of labor supply have much effect on hours, regardless of whether or not the job has changed. One must have variables that have a strong influence on desired hours to implement the approach used in the paper.

quit. In this case, our results would simply reflect heterogeneity of preferences and would neither imply that hours constraints are important within jobs nor that shifts in labor supply preferences induce individuals to change employers.

The second alternative is a model in which hours in a given job are determined by the employer but job changing costs and search costs are so low that workers can costlessly adjust hours by moving to a new employer following a change in preferences. If the labor supply response parameters are heterogenous, then individuals with the larger labor supply responses will be more likely to switch jobs to escape hours constraints. As a result, one would expect our finding that $\alpha + \beta$ exceeds α in absolute value. Under this scenario of employer determined hours with perfect mobility, changes in hours preferences have important implications for the analysis of job mobility but the hours choices of worker are unconstrained. From the point of view of labor supply analysis, there is no meaningful distinction between varying hours within and across firms.

Since the essence of both alternative interpretations of the results is that the labor supply parameters are related to the propensity to quit, we estimate models in which the hours responses to the labor supply variables depend upon estimates of the quit propensity. Our basic results are unchanged and indicate that workers have difficulty adjusting hours without changing employers. The evidence in the present study in combination with evidence from the job search literature and evidence from Altonji and Paxson (1986) that persons who switch jobs due to a layoff experience hours changes that are if anything larger than persons who switch jobs voluntarily indicates that individuals are unable to avoid hours constraints at low cost by changing jobs.

II. Data

The data are from the (1968–83) family/individuals PSID file. To be included, individuals must have been either a head of household or a wife in 1979, 1980, and 1981. Observations on wives for a particular year are used only if the individual was between the ages of 18 and 60 inclusive, was not retired, and was married during all of the previous four years. In the hours change equations, annual work hours had to be positive in both time periods over which the change in hours was computed.

The timing of the variables requires discussion. The surveys were conducted in the spring of each year (usually around March or April). The hours measures correspond to hours worked in the calendar year (January to January) before the survey. The quit indicator provided by the survey indicates whether a quit occurred in the year before the *survey* (i.e., March to March). The fact that hours refer to hours worked in the prior calendar year and the quit measure refers to the survey year poses a particular problem for the hours change equations. If a quit is reported to have occurred between March of $t - 1$ and March of t or between March of year t and March of year $t + 1$, hours in calendar year t might refer to hours on the old job, hours on the new job, or a mixture of both. To minimize this problem we measure hours changes (hours per week, weeks per year, and hours per year on the main job) over a four year interval: in terms of Equation (1), $m = 4$. The quit indicator Q_{it}^{m} is equal to 1 if a quit occurred in $t - 2$ or in $t - 3$. We measure quits in the middle of the interval to reduce the probability that the hours measures in t or in $t - 4$ are a mixture of hours in the new and old job.[4] The changes in demographic variables are changes between $t - 1$ and $t - 3$.[5] Due to limited information on the mobility of married women in the early years of the PSID, we can only use the 1983–79 and 1982–78 first-differences of hours in the hours change analysis. The hours level analysis in Table 1 uses observations on hours from 1970 to 1983, and information from 1968–83 was used to construct some of the other variables used in this analysis.

The most important variables in the analysis are those describing the composition of children in the family. We classify children according to whether they were born in the last two years, whether they are preschool (under age 6 but not newborn), and whether they are in school (age 6 to 17, inclusive.) The other labor supply determinants we focus on include changes in other family income (total family income excluding the wife's labor income), changes in the spouse's hours of unemployment, and changes in the spouse's health status.

The actual hours change equations estimated have a slightly different form from that specified by Equation (1). First, many of the labor supply determinants are discrete rather than continuous; it is inappropriate to

4. See Altonji and Paxson (1986) for a way to limit the sample to observations in which the hours measure refers to only one job. We do not use this approach here because it results in the elimination of large numbers of observations and biases the sample composition toward individuals who change jobs infrequently.

5. Some of the demographic variables used, such as total family income excluding the individual's labor income and hours worked, correspond to the calendar year (i.e., January to January) preceding the survey, and other variables, such as the spouse's health status and the number of children, refer to the actual time of the survey. Variables that refer to the survey date are lagged once, so that they will refer to the same time period as calendar year variables. For example, health status measured at the survey in year $t - 1$ is treated as contemporaneous with nonlabor income for calendar year $t - 1$ recorded in the year t survey.

simply compute changes in these variables. For changes in the spouse's health status, we construct two variables. The first equals 1 if the spouse gets over a health problem which affects his work ability (and 0 otherwise); the second equals 1 if the spouse acquires a health problem.

The treatment of the number and age composition of children is more complicated. First, there are many possible combinations of numbers of children in various age groups. Second, there may be interaction effects between changes in the composition of children and the current composition. For example, the effect on hours of a birth might depend on whether or not there are already other children at home. To account for these interactions, we construct three "transition" variables: the birth of a child, a child entering school, and a child finishing school. We then interact these transition variables with variables indicating the composition of other children in the household. For example, the variable "child enters school—some other preschoolers" indicates that a child entered school between $t - 3$ and $t - 1$, and that there were still other preschoolers in $t - 1$.[6]

Finally, we include a set of variables that may affect the average hours change, including age, education, and race. We do not control for wage rates because of measurement error in earnings divided by hours.[7] The coefficients for these variables are constrained to be the same for quitters and nonquitters. We do allow the intercept to vary depending on whether or not a quit occurred.[8]

The hours change equations are estimated with ordinary least squares. We present a basic set of results in Section III and discuss and attempt to deal with selection bias in Section IV. We report conventional OLS and White t-statistics. The White t-statistics, which tend to be smaller, account for serial correlation and heteroscedasticity but may be subject to larger sampling variation.

6. We assume that children enter school between the ages of five and six, and that children leave school between the ages of 17 and 18.

7. We estimated the models in Table 2 with the change in the log of earnings divided by hours controlled for. The wage coefficients were negative and the effects of the demographic variables are similar to those in Table 2. We also estimated the equations by *2SLS* using the change in a reported hourly wage rate (measured at the survey dates in t and $t - 4$) for 55 percent of the sample for which this wage rate is available. The wage coefficients are implausibly large for those who quit but are very imprecisely estimated. They are statistically significant only in the case of weeks worked per year. The differences between quits and nonquits in the responses to the demographic variables, however, do not change very much relative to standard errors.

8. We also estimated the models in Table 2 without imposing the restriction that control variables have the same coefficients for quits and nonquits. There is some evidence against the restrictions and a loss in precision in the estimates, but little change in the main results.

Table 1
Level Hours Equations (OLS t-statistics in parentheses)

	HOURS/WEEK	WEEKS/YEAR	HOURS/YEAR	MEAN (Standard Deviation)
	(1)	(2)	(3)	(4)
(1) Spouse has no disability in t, $t-1$, or $t-2$.	−.3165	.0596	−11.754	.8216
	(1.21)	(.18)	(.72)	
(2) Spouse has disability in t, $t-1$, or $t-2$.	.8537	−1.685	−20.867	.0586
	(2.05)	(3.18)	(.80)	
(3) $N=1, P=0, S=0$	−3.463	−5.987	−343.81	.0214
	(5.82)	(7.92)	(9.30)	
(4) $N=1, P=1, S=0$	−3.784	−5.556	−334.07	.0111
	(4.71)	(5.45)	(6.69)	
(5) $N=1, P>1, S=0$	−2.203	−3.405	−277.33	.0025
	(1.33)	(1.62)	(2.70)	
(6) $N=1, P=0, S>0$	−4.549	−7.828	−418.03	.0189
	(7.26)	(9.84)	(10.74)	
(7) $N=1, P=1, S>0$	−3.480	−7.027	−408.37	.0083
	(3.75)	(5.97)	(7.09)	

(8) $N = 1, P > 1, S > 0$	−4.273	−11.911	−539.28	.0023
	(2.46)	(5.40)	(5.00)	
(9) $N = 0, P = 1, S = 0$	−3.187	−2.760	−223.40	.0814
	(9.26)	(6.31)	(10.45)	
(10) $N = 0, P > 1, S = 0$	−4.964	−4.166	−336.12	.0257
	(9.07)	(5.99)	(9.89)	
(11) $N = 0, P = 0, S > 0$	−2.376	−2.307	−164.95	.3891
	(10.54)	(8.06)	(11.78)	
(12) $N = 0, P = 1, S > 0$	−3.192	−3.604	−246.07	.1030
	(9.88)	(8.78)	(12.26)	
(13) $N = 0, P > 1, S > 0$	−3.596	−5.171	−316.58	.0252
	(6.44)	(7.28)	(9.12)	
(14) Other income/1,000 in t	−.1286	−.1223	−8.397	10.35
	(10.18)	(7.62)	(10.69)	(7.20)
(15) Spouse's hours of	.0508	−.0865	−1.536	.7089
unemployment/100 in t	(1.49)	(2.00)	(.72)	(2.46)
R^2	.04	.04	.04	
Degrees of freedom	17,068	17,068	17,068	

Note: Means of dependent variables: *HOURS/WEEK* = 34.79 (10.96), *WEEKS/YEAR* = 39.37 (13.96), *HOURS/YEAR* = 1,396.86 (681.83). The variables in rows 3–13 reflect the composition of children. "N" refers to number of newborns in family (N = 0 or 1). "P" refers to number of children less than 6 years old, excluding newborns. "S" refers to number of children between 6 and 17, inclusive. Other variables included age, age^2, age^3, years of education, race and an intercept. The sample consists of wives who worked positive hours in year t.

III. Basic Results

In order to interpret the hours change results, one must know how each indicator of labor supply preferences is related to desired hours. For example, if women with ill husbands tend to work more, we expect that women whose husbands become ill will increase their hours, and that this increase in hours will be larger for those women who change jobs than for those who do not change jobs. To provide a rough guide to the effect of various labor supply variables on desired hours, we begin by discussing regressions of the *level* of hours on the set of labor supply preference variables and control variables used in the analysis of hours changes. Where possible we use several years of information on the labor supply characteristics to allow time for actual hours to have adjusted toward desired hours. The sample consists of women who worked positive hours in year t. The results for hours per week, weeks per year, and hours per year are in Columns 1, 2, and 3 (respectively) of Table 1.

The variable "spouse has no disability" is equal to one if the spouse had no illness that limited work in years t, $t - 1$, and $t - 2$. The variable "spouse has disability" equals one if the spouse had an illness that limited work in each of the three years. The omitted group consists of individuals whose husbands' health status changed at least once between t and $t - 2$. Neither variable has much effect on the hours measures, which suggests that they will not have a strong relationship to hours changes with or without quits.

We use a more elaborate specification for the effects of children than is typical in the literature.[9] The variables N, P, and S are dummy variables that indicate the presence of children of various ages in the household. N is equal to one if there are one or more newborns in the family in year t, and 0 otherwise. P equals the number of children less than six years old, excluding newborns. S equals the number of children between six and 17 inclusive. The reference group in the equation consists of wives who do not have children in year t. As might be expected, the results show that working wives who have a newborn child work considerably less than wives who do not have children. For example, working wives who have a newborn child and no other children (see Row 3) work 3.46 fewer hours per week and almost six fewer weeks per year than

9. See, for example, Jakubson (1988), who examines the effects of number of children less than 18 and number of children less than 6 in a life-cycle labor supply framework. He finds some evidence for a complicated dynamic relationship between children and measured work hours. Jakubson notes his specification of the children variables may be too restrictive, but also reports that he was unsuccessful in estimating a model with a more elaborate specification.

women without children. The results also suggest that additional children given the presence of a newborn child have much smaller effects on hours per week, weeks per year, and hours per year. For example, wives who have a newborn child and an additional preschool child work almost the same hours as wives with a newborn child only. Finally, the results suggest that women who have school-age children but no newborn work somewhat more than women with newborn children, and women with school-age children but neither newborns nor preschool children work more than women who have newborns or preschool children.

The results show a negative relationship between other income (which includes earned income of spouse) and the hours per week, weeks per year, and hours per year of the wife.

The strong relationship between the child composition variables and work hours implies that, if hours constraints do in fact restrict hours choices, we should find a strong relationship between the child composition variables and hours changes for quitters. We might also expect to see a negative relationship between changes in other income and changes in hours. The weak relationship between the hours level and spouse's health variables implies that changes in these variables should have little effect on hours changes for quitters and nonquitters.

As a crude check on the possibility that restricting our sample to women with positive hours in t and $t - 4$ colors the results, we have also estimated the level equation after restricting the sample to women who were in the sample in $t - 4$ (14,488 out of 17,688 person year observations) and who worked positive hours in $t - 4$ as well as t (10,575 person year observations). The results are basically similar to those in Table 1, and the coefficient on other family income is negative and statistically significant. When we estimate the level equation for hours using the hours in t for the sample used to estimate the hours change equations (i.e., $t = 1982$ or $t = 1983$), we also obtain similar results. The coefficient on nonlabor income, however, is positive and statistically significant rather than negative. We have no explanation for this anomaly other than sampling error.

A. *Results for Hours Changes*

To provide an overall feeling for the results given that we work with so many different labor supply variables, we first estimate the model:

$$(2) \quad \Delta^m H_{it} = \Delta^m X_{it} [\alpha + \alpha\phi Q_{it}^m] + \varepsilon_{it},$$

which imposes the restriction that the response to a change in X for those who do not quit is equal to $1/(1 + \phi)$ of the response of those who do quit.[10] [In terms of Equation (1), $\beta = \alpha\phi$.] The estimates (standard errors) of $1/(1 + \phi)$ are .328 (.080) for hours/week, .268 (.107) for weeks/year, and .335 (.083) for annual hours. In summary, the response of hours to the labor supply preference variables is only about one-third as large when an individual remains with the same employer.

The unrestricted equations for the change in hours between t and $t - 4$ are in Table 2. The coefficients on the labor supply variables for the change in hours per week, the change in weeks per year, and the change in hours per year are in Columns 1a, 2a, and 3a, respectively, for persons who did not quit their jobs during periods $t - 2$ or $t - 3$, and in Columns 1b, 2b, and 3b, respectively, for persons who did quit their jobs in $t - 2$ or $t - 3$. We have starred the coefficients for quits in Columns 1b, 2b, and 3b that are significantly different at the .05 level than the coefficients for nonquits in 1a, 2a, and 3a.

The most striking result is in Row 1, which shows the effect of a birth when the wife has no other children in $t - 1$. Hours per week fall by 5.86 hours for wives who did not quit their jobs and by 11.8 hours for wives who did quit their jobs. Similarly, weeks per year fall by 2.7 for wives who did not quit their jobs and by 8.4 for wives who did quit their jobs. A birth when no other children are present produces a much larger reduction in annual hours for those who quit (−735.6) than for those who did not quit (−362.2). Given that the sample means for hours per week, weeks per year and hours per year are 34.8, 39.4, and 1,396.9, respectively, these effects are quite large. The birth of a child between years $t - 3$ and $t - 1$ when no other children are present in $t - 1$ leads to a reduction in hours per year that is more than half of the mean value of hours per year.

The effects of a birth when there are other preschoolers in $t - 1$ but no school-age children are relatively small in magnitude and are not statistically significant (Row 2). Row 4 shows that a birth when there are other preschoolers and school-age children in $t - 1$ has a substantial negative effect on hours for those who change jobs. For example, hours per week fall by −2.8 hours for those who do not quit and by −7.2 hours for those who do quit. Weeks per year fall by −1.6 weeks for those who do not quit and −11.0 weeks for those who do. The coefficients for quits are statistically significant at the 10 percent level or higher using the OLS t-statistics. The results in Row 3 for the effect of a birth when there are no other preschoolers and some school-age children also show a substantial

10. This restriction (and all others like it in what follows) does not pass a likelihood ratio test. The purpose of the restriction is to provide a descriptive summary measure of the differences in hours responses between quitters and non-quitters.

negative effect for those who quit their jobs. (The effect on hours per week is actually positive for those who do not quit their jobs.)

We conclude that the occurrence of a birth, particularly the birth of the first child, has a strong negative effect on hours worked, and this effect is much larger for those who change jobs. In evaluating these results one should keep in mind that the *sample consists of wives who work positive annual hours in year* t *and year* t − *4*. Thus, the reductions in hours for those who quit do not reflect movements out of the labor force.

Row 5 shows the effect on hours of a child entering school when there are no other preschoolers in the household, and Row 6 shows the effect on hours of a child entering school when there are some other preschoolers. In both cases, a child entering school is positively related to hours changes for those who have quit their jobs. The estimates in Row 5 for weeks/year and hours/year show statistically significant differences between quitters and nonquitters. We are somewhat surprised to find that wives who had children who entered school but still had other preschoolers show a substantially larger increase in the hours measures than the wives who had a child who entered school and had no other preschoolers.

The results in Row 7 show the hours responses that occur when a child finishes school and there are some other children in the household. We find almost no effect for those who do not change jobs, and a large positive effect for those who do change jobs. The coefficients for the job changers are not precisely estimated, however, and only the effect for hours per year would be judged statistically significant based on the White t-statistics. The imprecision for this and some of the other child composition variables is due in part to the small number of observations on wives who quit and who fall within each of the categories for the child composition variables. (See the mean values for the child composition variables in the last two columns of Table 2). Despite this imprecision, the responses for quitters and nonquitters are significantly different for hours per week (using the OLS standard errors) and hours per year (using both OLS and White standard errors.)

Row 8 shows the effect on hours of a child finishing school when there are no other children in the household in year $t - 1$. Once again we find a much larger positive effect on hours for wives who have moved from one job to another than for wives who remained in the same job. Hours per week for job changers increased by 9.7 hours with the White t-statistic of 2.35. Hours per week for wives who do not change jobs increased by only .65. The estimates for weeks per year and hours per year for job changers, however, are not statistically significant and are mixed in sign.

Rows 9 and 10 show the effects of changes in the husband's disability status on hours changes for the wife, and Row 12 shows the effect of changes in the spouse's hours of unemployment. None of these coeffi-

Table 2
Hours Change Equations (OLS t-statistics in parentheses, White t-statistics in brackets)

	ΔHOURS/WEEK		ΔWEEKS/YEAR		ΔHOURS/YEAR		MEAN (Standard Deviation)	
	Nonquits	Quits	Nonquits	Quits	Nonquits	Quits	Nonquit	Quit
	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)		
(1) Birth, no other children	−5.859	−11.77	−2.703	−8.395	−362.19	−735.57	.037	.027
	(4.72)	(7.66)*	(1.60)	(4.01)*	(4.44)	(7.29)*		
	[3.48]	[5.20]*	[1.48]	[3.60]*	[3.40]	[5.76]*		
(2) Birth, other preschoolers, no school age children	.4518	.9283	3.485	.3204	122.75	72.89	.030	.013
	(.34)	(.45)	(1.91)	(.11)	(1.39)	(.53)		
	[.36]	[.35]	[1.90]	[.09]	[1.66]	[.45]		
(3) Birth, no other preschoolers, some school age children	3.529	−2.387	−2.234	−4.303	73.04	−242.84	.023	.009
	(2.28)	(.86)	(1.06)	(1.13)	(.72)	(1.33)		
	[1.84]	[1.23]*	[1.15]	[1.12]	[.77]	[1.39]		
(4) Birth, other preschoolers and school age children	−2.811	−6.215	−1.519	−11.04	−126.68	−534.94	.016	.005
	(1.55)	(1.70)	(.62)	(2.22)	(1.06)	(2.23)		
	[1.58]	[1.20]	[.51]	[1.31]	[.90]	[1.44]		
(5) Child enters school, some preschoolers	−1.339	1.997	2.088	11.2	6.611	431.17	.052	.014
	(1.25)	(.85)	(1.42)	(3.49)*	(.09)	(2.78)*		
	[1.32]	[.96]	[1.14]	[2.67]*	[.08]	[2.56]*		

(6) Child enters school,	1.407	.339	1.852	1.866	104.34	83.42	.047	.014
no preschoolers	(1.31)	(.17)	(1.27)	(.68)	(1.48)	(.63)		
	[1.24]	[.12]	[1.30]	[.44]	[1.48]	[.42]		
(7) Child finishes	−.0498	4.403	−.0639	4.092	−4.495	412.69	.089	.013
school, some	(.06)	(2.07)*	(.06)	(1.41)	(.08)	(2.96)*		
other children	[.07]	[1.55]	[.06]	[1.09]	[.08]	[2.65]*		
(8) Child finishes	.6540	9.705	.2001	−3.663	10.02	160.82	.046	.004
school, no	(.59)	(2.72)*	(.13)	(.75)	(.14)	(.69)		
other children	[.81]	[1.98]*	[.14]	[.48]	[.16]	[.43]		
(9) Spouse gets rid of a	.9068	2.155	−2.389	−2.356	6.298	−59.47	.029	.006
disability that	(.69)	(.74)	(1.33)	(.59)	(.07)	(.31)		
limits work	[.56]	[.62]	[1.34]	[.46]	[.06]	[.28]		
(10) Spouse acquires a	−1.514	3.757	.5603	−1.036	−23.66	79.28	.039	.012
disability that	(1.32)	(1.77)*	(.36)	(.36)	(.31)	(.57)		
limits work	[1.71]	[1.44]*	[.39]	[.27]	[.40]	[.41]		
(11) Δ (Other	.0025	.2686	.0491	.0967	2.164	14.60	−1.82	−.458
income)/1,000	(.07)	(3.23)*	(.98)	(.85)	(.89)	(2.67)*	(6.1)	(2.8)
	[.07]	[2.31]*	[.98]	[.71]	[.89]	[2.30]*		
(12) ΔSpouse's hours	−.0036	−.1444	.0243	.0523	1.583	1.771	.259	.047
unemployment/100	(.04)	(.94)	(.20)	(.25)	(.27)	(.18)	(2.5)	(1.5)
	[.07]	[.70]	[.18]	[.19]	[.29]	[.14]		
R^2	.078		.042		.070			
Degrees of freedom	1,973		1,973		1,973			

Note: A "*" by a t-statistic means that the coefficient for quits is significantly different from the coefficient for nonquits, at the 5 percent level or better. Other variables included: an intercept, age, age,^{2} and age^3, education and race. For all these variables except the intercept, coefficients were constrained to be the same for quits and nonquits. The sample is of wives with positive hours in t and $t - 4$. The exogenous variables refer to changes between $t - 1$ and $t - 3$. An observation is a quit if a quit occurred in $t - 2$ or $t - 3$. Means for dependent variables: Δ*HOURS/WEEK* = .3498 (10.13), Δ*WEEKS/YEAR* = 1.988 (13.55), Δ*HOURS/YEAR* = 83.16 (662.66).

cients are statistically significant; this is to be expected given evidence in Table 1 that the spouse's unemployment and the spouse's health status have relatively weak effects on hours levels.

The results in Row 11 for changes in other income appear to support the idea that hours-constraints are operative. We find that increases in other family income produce small but statistically significant increases in hours per week and hours per year for quitters, and have no effect on hours for nonquitters. Presumably, this result is related to the anomalous positive coefficients on other income that we obtain when we estimate the level equations in Table 1 on the sample used to estimate the hours change equations.[11]

IV. Accounting for Heterogeneity in Labor Supply and Job Mobility

In this section we first use a simple model of hours determination and job mobility to show that if (1) there is heterogeneity in the labor supply parameters G and mobility costs C differ across individuals and (2) persons with large (absolute) values of G tend to have low mobility costs, then the sample of quits will show larger hours responses than the sample of nonquits *even if all individuals are free to adjust their hours on the current job*. We then consider whether our results are also consistent with a model in which workers are not free to adjust hours on the current job but mobility costs are small and search costs are low in the sense that workers may easily locate a job offering the hours they wish to work. We conclude that if this "employer determined hours—perfect mobility model" is correct and labor supply preferences are heterogenous, then one would expect the results in Table 2 even though most workers can cheaply avoid hours constraints in particular jobs by changing employers.

We then present additional empirical estimates which deal with the potential problem of bias against the null hypothesis of no hours constraints within jobs by permitting the hours response to $\Delta^m X_{it}$ to depend upon an estimate of the quit probability for each individual.

Consider the following simple model of work hours and mobility.[12] Desired hours L_{it} and actual hours H_{ijt} of worker i in firm j in period t are determined by (3) and (4) below:

11. Controlling for the change in the wage rate did not eliminate the problem.

12. The issue of selection bias associated with heterogenous coefficients is certainly not unique to this paper. See, for example, Rosenzweig and Wolpin's (1989) analysis of migration.

(3) $L_{it} = X_{it} G_i + v_{it}$,

(4) $H_{ijt} = AL_{it} + (1 - A)F_{ij}$.

In the above equations:

X_{it} = vector of observed preference shifters in period t,

v_{it} = unobserved preference shifter in period t,

G_i = sensitivity of desired hours to X_{it},

F_{ij} = the number of hours firm j would prefer i to work,

A = weight placed on desired labor supply in hours determination.

We normalize the elements of X so that the vector $G_i > 0$. G_i varies across i but is always nonnegative. In (3) we implicitly assume that desired hours are independent of job characteristics. This is a reasonable approximation if either the sensitivity of labor supply to wages and working conditions is small, or the variation in wages and working conditions affecting labor supply is small.

Individuals have static expectations about X_{it} and v_{it}. The utility of job j for person i is:

(5) $V_{ijt} = Z_{ij} - |L_{it} - H_{ijt}|$,

where Z_{ij} is an index summarizing the effects of wages and other job characteristics on the utility from the job. We normalize Z and V so that $|L_{it} - H_{ijt}|$ and Z_{ij} have unit coefficients in (5).

The utility of an alternative job with firm j' that has characteristics $Z_{ij'}$ and an hours demand of $F_{ij'}$ is:

(6) $V_{ij't} = Z_{ij'} - |L_{it} - H_{ij't}|$.

Conditional upon having offer j' the individual changes jobs if $V_{ij't}$ exceeds the utility V_{ijt} of the current job by more than mobility costs C_i:

(7) Quit if: $V_{ij't} - V_{ijt} > C_i$.

To keep the example simple, suppose that the individual happens to have located an opening in which the $F_{ij'} = L_{it}$. (Note that under the null hypothesis that workers can freely choose hours, $L_{it} = H_{ij't}$ in all jobs.) Also, rewrite L_{it} as $L_{it} - L_{it-m} + L_{it-m}$ and substitute for L and H in (6) using (3), (4), and (5). Then (7) becomes:

(8) $(1 - A)|\Delta^m X_{it} G_i + \Delta^m v_{it} + L_{it-m} - F_{ij}| + (Z_{ij'} - Z_{ij}) > C_i$.

The first term in (8) is the gap between desired and actual hours on job j. The term $(Z_{ij'} - Z_{ij})$ is the value that i places on other characteristics of job j and j', including wages. The equation implies that persons with low values of C_i tend to have higher quit probabilities.

Now consider the implications of the above analysis for the estimation of β in (1) under the null hypothesis that individuals may freely choose hours, with $A = 1$. In this case hours considerations drop out of (8) and mobility is driven only by comparisons of $(Z_{ij'} - Z_{ij})$ with C_i. If C_i and the elements of G_i happen to be negatively correlated, however, then individuals who quit tend to have both lower C_i and higher G_i than those who do not. Consequently, in this case the response of work hours to a change in X is larger for those who quit than for those who do not.[13]

Since the essence of the problem is that the quit propensity P_i and G_i are related, one way to eliminate or at least reduce the bias is to explicitly account for dependence between G_i and P_i. One may decompose G_i into its expectation given P_i and an orthogonal error. Assume the relationship is linear. Then

$$(9)\quad G_i = G + P_i\psi + g_i,$$

where g_i is uncorrelated with P_i and $G + P_i\psi$ is the expectation of G_i conditional on P_i. Normalize P_i and g_i to have mean 0 so that G is the mean of G_i.

Using (3), (4), and (9) the hours change equation is:

$$(10)\quad \Delta^m H_{ijt} = \Delta^m X_{it}\ A[G + P_i\psi + g_i] + A\Delta^m v_{it} + (1 - A)\,\Delta^m F_{ijt},$$

where $\Delta^m F_{ijt} = 0$ if the job does not change and $F_{ij'} - F_{ij}$ if a quit occurs. Under the null hypothesis that workers may freely choose hours on any given job (A = 1), (10) reduces to:

$$(11)\quad \Delta^m H_{ijt} = \Delta^m X_{it}\, G + \Delta^m X_{it} P_i\psi + \varepsilon_{it},$$

where $\varepsilon_{it} = \Delta^m v_{it} + \Delta^m X_{it} g_i$.

Equation (11) says that hours changes should be the same for quits and nonquits, and to test this we estimate:

$$(12)\quad \Delta^m H_{ijt} = \Delta^m X_{it}\,[G + P_i\psi + Q_{it}^m\,\beta] + \varepsilon_{it},$$

where β is 0 under the null hypothesis that hours are flexible. In estimating (12) we include P_i and Q_{it}^m as controls as well as the other control variables used in the basic results.

Although β is 0 under the null, one must consider the possibility that correlation between Q and the error term $\Delta^m X_{it}\ g_i + \Delta^m v_{it}$ will bias β

13. We do not know of any evidence on the correlation between C and G_i. Below we find that G_i is positively related to the quit probability. A priori, one might argue that persons with highly variable labor supply preferences are likely to withdraw from the labor force entirely at times and as a result may tend to choose jobs with little deferred compensation.

away from 0. Q_{it}^m is uncorrelated with g_i by definition of g_i in (9) and the fact that the error relating Q_{it}^m to the quit probability P_i has mean 0 for each i. $\Delta^m X_{it}\, g_i + \Delta^m v_{it}$ is uncorrelated with $\Delta^m X_{it}$ or Q_{it}^m because Q_{it}^m and $\Delta^m X_{it}$ are both included in (12). Unfortunately, these orthogonality conditions (without statistical independence) are not enough to guarantee that nonlinear functions of Q_{it}^m and $\Delta^m X_{it}$, such as the product $\Delta^m X_{it}\, Q_{it}^m$, are uncorrelated with $(\Delta^m X_{it}\, g_i + \Delta^m v_{it})$. We proceed with (12) under the assumption that any bias associated with the nonlinearity is smaller than the bias associated with ignoring the heterogeneity in P_i.

A. *Implications of the "Employer Determined Hours—Perfect Mobility Model" when Labor Supply is Heterogenous*

Since a key question is whether individuals can costlessly adjust work hours, either on their current job or another job, consider the case in which A is less than 1 but mobility costs are small and the same for everyone, and search costs are sufficiently low that individuals may easily locate a job requiring L_{it} hours without having to settle for a less desirable value of Z. Assuming that $\Delta^m X_{it}$ is largely correlated with the other terms that affect quits, (10) implies that individuals with large values of G_i will be more likely to quit in response to change in X. As a result the expectation of β in (1) is nonzero.

If hours may only be partially adjusted on the current job, workers seek jobs that are somewhat closer to their desired hours given the change in X, which induces a correlation between $\Delta^m F_{ijt}$ and $\Delta^m X_{it}$. For those who quit let the least squares projection of $\Delta^m F_{ijt}$ on $\Delta^m X_{it}$, P_i, and $\Delta^m X_{it} P_i$ be:

$$(13) \quad \Delta^m F_{ijt} = [\Delta^m X_{it}\beta + \Delta^m X_{it} P_i \beta_1 + P_i \beta_2]/(1 - A) + f_{it}/(1 - A).$$

where $\beta/(1 - A)$, $\beta_1/(1 - A)$, and $\beta_2/(1 - A)$ are the parameters of the projection and $f_{it}/(1 - A)$ as an orthogonal error component. After substitution the hours change equation is:

$$(14) \quad \Delta^m H_{ijt} = \Delta^m X_{it}\,[AG + P_i A\psi + Q_{it}^m \beta + P_i Q_{it}^m \beta_1] + \varepsilon_{it},$$

where $\varepsilon_{it} = \Delta^m X_{it}\, A g_i + A\Delta^m v_{it} + Q_{it}^m f_{it}$. Controls for $P_i Q_{it}^m$, P_i, and Q_{it}^m are also included.

By examining the elements of $\beta + P_i\beta_1$ for particular values of P_i, we can examine whether hours responses to labor supply shifters are larger for quitters, controlling for the fact that the size of the labor supply shifts may be systematically related to the quit probability. The coefficients of (13) and therefore (14) will be influenced by the ease with which individuals, conditional on the decision to quit, are able to move to a job offering L_{it} hours. While we can obtain information about the extent to which individuals can adjust hours without changing jobs by examining $\beta + P_i\beta_1$, we cannot determine the ease with which they are able to move in response to a change in desired hours from (14) alone. However, other evidence runs counter to the assumption of near perfect mobility.

First, the significant amount of time that individuals spend searching for jobs and low turnover rates among individuals with more than a year or two of tenure suggest that mobility and search costs are substantial. The evidence that the wages of a given worker vary substantially from one employer to another is hard to reconcile with a labor market in which information is cheap and mobility costs are low.[14] Indeed, the basic premise of the huge literature on labor market search is that search and mobility costs are substantial.[15]

Second, the evidence in Altonji and Paxson (1986) also is inconsistent with the perfect mobility story.[16] In that paper we show that a large fraction of the variance in hours work is job specific. Individuals experience much larger changes in hours when they change employers than when they remain with the same firm. One interpretation of this finding is that the hours requirements of a job have a significant influence on job choice. We then compare the "employer determined hours-perfect mobility model and the "employer determined hours-imperfect mobility model" taking into account the possibility that the large hours changes associated with quits are due to the fact that individuals quit to adjust hours in response to preferences changes. We find the variance in hours changes for those who experience a layoff is as large or larger than the variance of those who quit. Assuming that the probability of a layoff is unrelated to heterogeneity across individuals in the change in desired hours, the fact that those who separate due to a layoff are no more likely to return to a job offering the same hours level as their previous one suggests workers cannot easily optimize work hours through job choice.

B. Estimation

To implement (12) or (14) we need a proxy for P_i. The spirit of the analysis is that there is fixed individual heterogeneity in mobility behavior. Presumably, this is at least partially revealed in past turnover behavior.

14. See, for example, Altonji and Shakotko (1987) or Topel (1991).

15. See Divine and Kiefer (1991) for a recent survey.

16. In Altonji and Paxson (1988) we find that overemployment and underemployment on the initial job and on the new job affects the relation between hours changes and wages changes for those who quit. The results suggest the mobility and search costs are such that workers must trade off the desirability of work hours in a particular job against wages and other job characteristics.

Consequently, we estimate P_i as the implied quit probability from a probit equation for Q_{it}^m in which the key explanatory variables are employer tenure and tenure squared in $t - m$. We also control for $\Delta^m X_{it}$ and for whether the individual changed the state or county of residence between t and $t - m$.[17] To the extent that P_i depends on variables that are left out of the quit model, some additional bias may arise under the null hypothesis. Also, to the extent that the labor supply coefficients G vary across t as well as across i, then controlling for P_i may be inadequate.

C. Results

To provide an overall assessment of the difference in hours response to labor supply changes for those who quit during period $t - 2$ or $t - 3$ and those who do not, we first estimate (12) with the restriction that $\beta = \phi G$ imposed on vectors G and β. This restriction implies that the response of nonquitters to a change in preferences relative to the response of quitters equals $1/(1 + \phi)$. The estimates (standard errors) of $1/(1 + \phi)$ are .348 (.091) for hours per week, .447 (.152) for weeks per year, and .391 (.101) for annual hours. These estimates are not substantially higher than the corresponding estimates that exclude P from the model. This implies that heterogeneity in labor supply parameters which happens to be associated with the quit probability is not solely responsible for our finding that hours adjustments in response to labor supply preferences are larger when individuals switch employers.[18]

We do, however, find evidence that people with higher quit probabilities have larger hours responses to changes in labor supply preferences independent of whether there is a quit. We estimate variants of Equation (12) which include the restrictions that $\beta = \phi G$ and $G = \phi_1\psi$, implying that, in percentage terms, the effect of an increase in the quit probability on the responsiveness of hours to changes in preferences is the same for

17. When estimating the restricted and unrestricted versions of (12) discussed below we also experimented with evaluating $\Delta^m X_{it}$ and the dummies for whether the individual changed residence at the sample means when forming the estimate of P_i. In this case we are relying solely on past mobility as summarized by job tenure to identify P_i. This made little difference in the results.

18. If *jobs* differ with respect to hours flexibility (i.e., the parameter A varies), then one would expect persons in jobs offering flexible hours to be less likely to quit in response to the change in labor supply preferences and to make larger hours adjustments than would be possible for persons in a representative job. As a result, the estimated difference between those who quit and those who do not quit in the effects of the labor supply determinants on hours may be understated. We investigated this issue by estimating models that included interactions among the labor supply preference variables and whether the individual was free to increase hours on the initial job, and found that this made little difference.

all variables in X. Estimates of ϕ_1 are .190 (.082) for hours/week, .097 (.035) for weeks per year, and .215 (.103) for hours/year. Estimates of $1/(1 + \phi)$ are .250 (.078), .489 (.213) and .314 (.092). The estimates of ϕ_1 indicate that people with lower quit probabilities do have smaller hours responses to changes in labor supply preferences than those with larger quit probabilities, regardless of whether a quit occurs. The estimates of $1/(1 + \phi)$ imply, however, that this heterogeneity cannot fully explain the differences in hours changes between those who do and do not quit.

We have also computed unrestricted estimates of Equation (12) and computed for nonquits and quits the effects of a change in X on the change in hours evaluated at the sample mean for the quit probability. [In terms of Equation (12), these effects are $G + \psi P$ for nonquits and $G + \beta + \psi P$ for quits, where P is the mean quit probability of .184.] The estimates are reported in Altonji and Paxson (1990) but are omitted here to save space. Without going into the details, the difference in responses for those who quit and those who do not are usually consistent with those reported Table 2, but the magnitude of the difference is reduced somewhat by controlling for P_i. For example, the reduction in hours associated with a birth when there were no other children is 5.55 hours larger if a quit occurs. When one does not control for P_i (Table 2), the corresponding estimate if 5.91. Evidently, heterogeneity in labor supply parameters is only a partial explanation for the finding that the hours response to labor supply preferences is larger if the person changes employers.

We have also estimated restricted and unrestricted versions of (14). We impose the restriction that at the sample mean of P_i the response to $\Delta^m X_{it}$ for those who do not quit is $1/(1 + \phi)$ of the response for quits. In terms of (14) the restriction is that $(\beta + P\beta_1) = \phi\,(AG + PA\psi)$, where P is the sample mean of the quit probability. The estimate (standard error) of $1/(1 + \phi)$ is .374 (.122) for hours per week, .570 (.198) for weeks per year, and .500 (.143) for annual hours. These estimates are somewhat larger than the estimates of $1/(1 + \phi)$ when the quit probability interactions are not included. They still indicate, however, that hours changes are much smaller for those who do not quit than for those who do.

The unrestricted estimates of Equation (14), presented in Altonji and Paxson (1990), also show larger hours changes for those who quit, although the differences between those who do and do not quit are somewhat smaller than those indicated by the model with no controls for the quit probability. For example, we find that a first birth produces a decline in hours per week that is 4.8 hours greater for those who quit (evaluated at the mean quit probability), in contrast to the difference of 5.91 implied by results of Table 2.

In general, the estimates of Equations (12) and (14) indicate that heterogeneity in preferences accounts for some of the differences between the hours changes of quitters and nonquitters. Even after controlling for heterogeneity, however, we still find that the response of hours to changes in preferences is much larger for those who quit than for those who do not. These results are inconsistent with the hypothesis of no hours constraints within jobs. Our results, when considered in conjunction with other evidence, are also inconsistent with the hypothesis that individuals are able to costlessly avoid hours constraints by changing jobs.

IV. Conclusion

Our main finding is that the effect of changes in the demographic structure of the family on wives' work hours are generally much larger for wives who change employers than for those who do not. This finding, which does not appear to be driven by heterogeneity in preferences, is consistent with the view that constraints on hours choice within individual firms limit the extent to which workers may change hours within a job following a change in labor supply preferences. Job changes following shifts in labor supply preferences may provide the opportunity to reduce any discrepancies between desired hours and actual hours.

Future research should replicate our analysis on another data set, such as the NLS, which contains better information about hours worked and reasons for leaving specific jobs than the PSID. The next major step is the estimation of a structural model of labor supply, hours constraints, and job mobility, since, as we have emphasized in the paper, changes in preferences may also affect the probability of a quit if jobs do consist of hours-wage packages. Unfortunately, estimation of even a simple structural model would be a formidable task and would require better data than we have.

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