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Joseph G. Altonji

Christina H. Paxson

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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.

Joseph G. Altonji
Department of Economics
Northwestern University
Evanston, IL 60208

Christina H. Paxson
Woodrow Wilson School
Princeton University
Princeton, NJ 08544

## I. Introduction

Most of the literature on labor supply is based on the assumption that workers can freely choose hours at a parametric wage. 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 then when it does not change.

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

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

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 $Q^m_{it}$ 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 $\beta$ is close to 0 and $\alpha$ is non-zero would suggest that the effect of preference changes on hours is independent of whether or not the job changed, implying hours flexibility within jobs. A finding that $\alpha$ is close to 0 (and $\beta$ non-zero) 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 $\alpha$ and $\beta$ will be of the same sign, but that $\alpha + \beta$ will be larger (in absolute value) than $\alpha$.

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 2.

Our basic result in Section 3 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.

However, there are two alternative interpretations of the basic results that we address theoretically and empirically in Section 4 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 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 $\alpha$ 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 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. However, other evidence must be considered in conjunction with the evidence in the present study if one is to rule out the possibility that individuals are able to avoid hours constraints at low cost by changing jobs.

## II. Data

The data are from the (1968-1983) 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 4 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 4 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 1983-1979 first difference 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 1968-1983 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 pre-school (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 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 effect the average hours change, including age, education and race. The coefficients for these variables are constrained to be the same for quitters and non-quitters. We do allow the intercept to vary depending on whether or not a quit occurred.

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 heteroscedasticity but may be subject to larger sampling variation.

## 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.[7] 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 seventeen, 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, 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 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 the these variables should have little effect on hours changes for quitters and non-quitters.

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:

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

which imposes the restriction that the response to X for those who do not quit is equal $1/(1+\phi)$ of the response of those who do quit.[8] (In terms of equation (1), $\beta = \alpha\phi$.) The estimates (OLS standard errors) of $1/(1+\phi)$ are .328 (.080) for hours/week, .268 (.107) for weeks/year, and .335 (.083) for annual hours. That is, the response of hours to the labor supply preference variables is only about 1/3 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 non-quits 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 (-358.5). Given that the sample means for hours per week, weeks per year and hours per year are 34.8, 39.4, and 1396.9, 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 effect 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 falls by -1.6 weeks for those who do not quit and -11.5 weeks for those who do. The coefficients for quits are statistically

significant at the 10% 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 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 non-quitters. 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. However, the coefficients for the job changers are not precisely estimated, 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 2 columns of Table 2). Despite this imprecision, the responses for quitters and non-quitters are significantly different for hours per week and hours per year.

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. However, the estimates for weeks per year and hours per year for job changers 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 coefficients 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 non-quitters. However, the results for the hours level equations in Table 1 suggest that increases in other family income should lead to hours reductions rather than hours increases. We do not have an explanation for the inconsistency in the results across the two tables.

**IV. Accounting for Heterogeneity in Labor Supply and Job Mobility.**

In the introduction we pointed out that heterogeneity in the responsiveness of desired hours to changes in labor supply determinants may bias the results against the hypothesis that individuals can adjust hours within jobs. 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 non-quits *even if all individual 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 in addition 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. Our results are inconsistent with the hypothesis of no hours constraints within jobs, although they do imply that part of the difference between the hours responses of quitters and non-quitters in Table 2 may be due to heterogeneity. While our results are not sufficient to rule out the hypothesis that individuals are able to costlessly avoid hours constraints by changing jobs, we point out other evidence that runs counter to this interpretation.

Consider the following simple model of work hours and mobility. Desired

hours $L_{it}$ and actual hours $H_{it}$ of worker i in firm j in period t are determined by (3) and (4) below:

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

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

In the above equations:

$X_{it}$ = vector of observed preferences 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 non-negative. In (3) we implicitly assume that desired hours is 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) $V_{ijt} = Z_{ij} - |L_{it} - H_{ij}|$,

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 $|L_{it} - H_{it}|$ and $Z_{ij}$ have unit coefficients in (5).

The utility of an alternative job with firm j' that has characteristics $Z_{ij'}$ and

hours demand $F_{ij'}$ is:

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

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 $F_{ij'} = L_{it}$. (This assumption is not restrictive under the null hypothesis that workers can freely choose hours, in which case $L_{it} = H_{ij'}$ in all jobs.) Also, re-write $L_{it}$ 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. A quit occurs if the sum of these terms exceeds mobility costs $C_i$. Consequently, persons with low values of $C_i$ have higher quit probabilities. The same argument could be made for any given value of $L_{it} - F_{ij'}$.

Now consider the implications of the above analysis for the estimation of $\beta$ 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 $C_i$. However, if $C_i$ and the elements of $G_i$ happen to be

negatively correlated, 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.[9]

Since the essence of the problem is that the quit propensity 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 the quit probability $P_i$. One may decompose $G_i$ into its expectation given $P_i$ and an orthogonal error. Assume the relationship is linear. (The linearity assumption could be relaxed, although we do not do so.)

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

where $g_i$ is uncorrelated with $P_i$ and $G + \psi P_i$ 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:

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

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:

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

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 non-quits, and to test this we estimate:

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

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

Although $\beta$ is 0 under the null, one must consider the possibility that correlation between Q and the error term $\Delta^m X_{it} Ag_i + \Delta^m v_{it}$ will bias $\beta$ away from 0. $Q^m_{it}$ is uncorrelated with $g_i$ by definition of $g_i$ in (9) and the fact that the error relating $Q^m_{it}$ to the quit probability $P_i$ has mean 0 for each i. $\Delta^m X_{it} Ag_i + \Delta^m v_{it}$ are both uncorrelated with $\Delta^m X_{it}$ and $Q^m_{it}$ because $Q^m_{it}$ 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 $Q^m_{it}$ and $\Delta^m X_{it}$, such as the product $\Delta^m X_{it} Q^m_{it}$, are uncorrelated with $(\Delta^m X_{it} g_i + \Delta^m v_{it})$. We believe (but have no formal proof or evidence) that any bias associated with the nonlinearity is likely to be second order and smaller than the bias in which we do not control for $P_i$ in the analysis.

Implications of Analysis for 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 (a) mobility costs are small and the same for everyone, and (b) 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 uncorrelated 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 $\beta$ in (1) is non-zero.

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:

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

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

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

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

By examining the elements of $\beta + P_i \beta_1$ for particular values of the quit probability P, we can examine whether hours responses to labor supply shifts responses 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, find it optimal 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 the perfect mobility model. First and foremost, 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.[10] Indeed, the basic premise of the huge literature on labor market search is that search and mobility costs are substantial.[11] Second, evidence in Altonji and Paxson (1988 and 1986) also is inconsistent with the perfect mobility story.[12]

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. Consequently, we estimate $P_i$ as the implied quit probability from a probit equation for $Q^m_{it}$ 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.[13] 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.

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 $\beta$. This restriction implies that the response of non-quitters 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 individual switch employers.[14]

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 all variables in X. Estimates of $\phi_1$ 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 $\phi_1$ 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. However, the estimates of $1/(1+\phi)$ imply that this heterogeneity cannot fully explain the differences in hours changes between those who do and do not quit.

In Table 3 we present the unrestricted estimates of equation (12). The coefficients under the columns for non-quits and quits are 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), we report $G + \psi P$ for non-quits and $G + \beta + \psi P$ for quits, where P is the mean quit probability of .18373. Also presented are the $\psi$ of equation (12). 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 in reduced somewhat by controlling for $P_i$. For example, Table 3 indicates that 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 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 larger for those who quit than for those who do not.

The unrestricted estimates of equation (14), presented in Table 4, 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, Table 4 implies that the effect of 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 non-quitters. However, even after controlling for heterogeneity 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.

## 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.[15]

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## Endnotes

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 Moffit (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), Moffit (1984), Lundberg (1984), and Biddle and Zarkin (1986)) 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.

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.

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 non-labor income for calendar year t-1 recorded in the year t survey.

6. We assume that children enter school between the ages of 5 and 6, and that children leave school between the ages of 17 and 18.

7. See for example, Jakubson (1986), 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.

8. 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.

9. We do not know of any evidence on the correlation between $C_i$ 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.

10. See, for example, Altonji and Shakotko (1987) or Topel (1990).

11. See Divine and Kiefer (1990) for a recent survey.

12. 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. In Altonji and Paxson (1986) 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 as the variance for 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 cannot easily optimize work hours through job choice.

13. 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.

14. If *jobs* differ with respect to hours flexibility (ie., 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.

15. In the course of doing this research, we formulated a simple model of job mobility and hours when jobs consist of hours-wage packages, and derived a likelihood function for quits, hours and wages conditional on changes in preferences. Notes on this model are available from the authors. To estimate the model, one would need information not only on determinants of labor supply, but also on the hours-wage distribution of offers which workers face: the quit probability is not just a function of the worker's degree of dissatisfaction with current hours, but also a function of the wage-hour offers the worker receives. Handling unobserved heterogeneity in labor supply preferences and in the wage offer distribution is a formidable task. For example, changes in preferences could shift desired hours to a level at which few job openings are available, thereby decreasing the quit probability. We have been able to show that, on average, the quit probability for people who have recently experienced a change in preferences will be higher than the quit probability for those whose preferences have been stable over time. However, this need not be true in any given case. We have not been successful in developing a dynamic model that we feel is sufficiently realistic to be worth taking to the data.

Estimation of a joint model of quits and hours changes is further complicated by the fact that the distribution of hours-wage offers may be itself a function of factors affecting desired labor supply. For example, suppose that the birth of a child reduces a woman's desired hours. If hours in her current job are not flexible, once might expect that the probability of a quit (to another job) will increase. On the other hand, the birth of a child may temporarily lower productivity. If the current employer is under an implicit contract not to reduce wages in response, this would lower the probability that a job offering an better hours-wage package can be found.

**TABLE 1**
**LEVEL HOURS EQUATIONS**
OLS - t-statistics in parentheses

| | HOURS/WEEK | WEEKS/YEAR | HOURS/YEAR | VARIABLE MEAN |
|---|---|---|---|---|
| | (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/1000 in t | -.1286 | -.1223 | -8.397 | 10.35 |
| | (10.18) | (7.62) | (10.69) | (7.20) |
| (15) SPOUSE'S HOURS OF UNEMPLOYMENT/100 in t | .0508 | -.0865 | -1.536 | .7089 |
| | (1.49) | (2.00) | (.72) | (2.46) |
| $R^2$ | .04 | .04 | .04 | |
| Degrees of Freedom | 17068 | 17068 | 17068 | |

Note: Means of dependent variables: HOURS/WEEK = 34.79 (10.96), WEEKS/YEAR = 39.37 (13.96), HOURS/YEAR = 1396.86 (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.

TABLE 2
HOURS CHANGE EQUATIONS
OLS t-statistics in parentheses, White t-statistics in brackets

| | ΔHOURS/WEEK | | ΔWEEKS/YEAR | | ΔHOURS/YEAR | | MEAN | (STD) |
|---|---|---|---|---|---|---|---|---|
| | NonQuits | Quits | NonQuits | Quits | NonQuits | Quits | NonQuit | Quit |
| | (1a) | (1b) | (2a) | (2b) | (3a) | (3b) | | |
| (1) BIRTH, NO OTHER | -5.8590 | -11.77 | -2.703 | -8.395 | -362.19 | -735.57 | .037 | .027 |
| CHILDREN | (4.72) | (7.66)* | (1.60) | (4.01)* | (4.44) | (7.29)* | | |
| | [4.17] | [5.67]* | [1.48] | [3.65]* | [3.82] | [6.13]* | | |
| (2) BIRTH, OTHER | .4518 | .9283 | 3.485 | .3204 | 122.75 | 72.89 | .030 | .013 |
| PRESCHOOLERS, NO | (.34) | (.45) | (1.91) | (.11) | (1.39) | (.53) | | |
| SCHOOL AGE CHILDREN | [.37] | [.38] | [1.94] | [.10] | [1.73] | [.48] | | |
| (3) BIRTH, NO OTHER | 3.529 | -2.387 | -2.234 | -4.303 | 73.04 | -242.84 | .023 | .009 |
| PRESCHOOLERS, SOME | (2.28) | (.86) | (1.06) | (1.13) | (.72) | (1.33) | | |
| SCHOOL AGE CHILDREN | [1.96] | [1.11]* | [1.15] | [1.09] | [.83] | [1.32]* | | |
| (4) BIRTH, OTHER | -2.811 | -6.215 | -1.519 | -11.04 | -126.68 | -534.94 | .016 | .005 |
| PRESCHOOLERS AND | (1.55) | (1.70) | (.62) | (2.22) | (1.06) | (2.23) | | |
| SCHOOL AGE CHILDREN | [1.6] | [1.19] | [.51] | [1.27] | [.91] | [1.43] | | |
| (5) CHILD ENTERS | -1.339 | 1.997 | 2.088 | 11.22 | 6.611 | 431.17 | .052 | .014 |
| SCHOOL, SOME | (1.25) | (.85) | (1.42) | (3.49)* | (.09) | (2.78)* | | |
| PRESCHOOLERS | [1.35] | [.93] | [1.18] | [2.62]* | [.08] | [2.43]* | | |
| (6) CHILD ENTERS | 1.407 | .339 | 1.852 | 1.866 | 104.34 | 83.42 | .047 | .014 |
| SCHOOL, NO | (1.31) | (.17) | (1.27) | (.68) | (1.48) | (.63) | | |
| PRESCHOOLERS | [1.29] | [.13] | [1.31] | [.45] | [1.53] | [.44] | | |
| (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] | [1.70]* | [.06] | [1.23] | [.09] | [2.81]* | | |
| (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 | [.79] | [2.35]* | [.15] | [.50] | [.17] | [.47] | | |
| (9) SPOUSE GETS RID | .9068 | 2.155 | -2.389 | -2.356 | 6.298 | -59.47 | .029 | .006 |
| OF A DISABILITY | (.69) | (.74) | (1.33) | (.59) | (.07) | (.31) | | |
| THAT LIMITS WORK | [.55] | [.62] | [1.29] | [.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] | [.28] | [.40] | [.41] | | |
| (11) Δ(OTHER | .0025 | .2686 | .0491 | .0967 | 2.164 | 14.60 | -1.82 | -.458 |
| INCOME)/1000 | (.07) | (3.23)* | (.98) | (.85) | (.89) | (2.67)* | (6.1) | (2.8) |
| | [.06] | [2.21]* | [.97] | [.70] | [.87] | [2.21]* | | |
| (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) |
| | [.06] | [.7] | [.18] | [.20] | [.28] | [.14] | | |
| $R^2$ | .078 | | .042 | | .070 | | | |
| Degrees of Freedom | 1973 | | 1973 | | 1973 | | | |

Note: A "*" by a t-statistic means that the coefficient for quits is significantly different from the coefficient for non-quits, at the 5% 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 non-quits. 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: ΔHOURS/WEEK=.3498 (10.13), ΔWEEKS/YEAR=1.988 (13.55), ΔHOURS/YEAR=83.16 (662.66).

**TABLE 3**
**HOURS CHANGE EQUATIONS WITH QUIT PROBABILITIES**
(Estimates of equation (12))

| | ΔHOURS/WEEK | | | ΔWEEKS/YEAR | | | ΔHOURS/YEAR | | |
|---|---|---|---|---|---|---|---|---|---|
| | NonQuit | Quit | P(Quit) | NonQuit | Quit | P(Quit) | NonQuit | Quit | P(Quit) |
| BIRTH, NO | -4.854 | -10.4 | -3.977 | -3.594 | -7.504 | -10.91 | -341.5 | -628.3 | -701.3 |
| OTHER CHILDREN | (2.86) | (4.76)* | (0.65) | (1.57) | (2.54) | (1.32) | (3.08) | (4.40)* | (1.75) |
| | [2.38] | [3.60]* | [0.48] | [1.76] | [2.55]* | [1.48] | [2.89] | [3.76]* | [1.57] |
| BIRTH, OTHER | 1.127 | 1.771 | -5.263 | 2.891 | 0.4167 | -5.042 | 136.3 | 137.6 | -587.0 |
| PRESCHOOLERS, NO | (0.74) | (0.71) | (0.66) | (1.41) | (0.12) | (0.47) | (1.37) | (0.85) | (1.12) |
| SCHOOL AGE CHILDREN | [0.91] | [0.66] | [0.91] | [1.55] | [0.12] | [0.55] | [1.84] | [0.83] | [1.73] |
| BIRTH, NO OTHER | 1.543 | -5.232 | 32.71 | -2.987 | -3.55 | -16.99 | -25.27 | -310.0 | 547.9 |
| PRESCHOOLERS, SOME | (0.80) | (1.71)* | (2.18) | (1.15) | (0.86) | (0.84) | (0.2) | (1.55) | (0.56) |
| SCHOOL AGE CHILDREN | [0.97] | [2.12]* | [2.14] | [1.48] | [0.89] | [0.92] | [0.27] | [1.68]* | [0.77] |
| BIRTH, OTHER | -3.598 | -8.676 | 32.11 | -2.497 | -9.952 | -14.25 | -171.9 | -546.2 | 157.2 |
| PRESCHOOLERS AND | (1.90) | (2.24) | (2.05) | (0.98) | (1.90) | (0.67) | (1.39) | (2.16) | (0.15) |
| SCHOOL AGE CHILDREN | [1.74] | [1.67] | [1.92] | [0.85] | [3.65] | [0.47] | [1.19] | [1.58] | [0.12] |
| CHILD TO SCHOOL, | -1.454 | 2.835 | -11.36 | 2.298 | 11.07 | 13.92 | 16.56 | 468.8 | -148.5 |
| NO PRESCHOOLERS | (1.35) | (1.13) | (0.98) | (1.57) | (3.27)* | (0.89) | (0.24) | (2.87)* | (0.2) |
| | [1.49] | [1.26]* | [0.98] | [1.33] | [2.51]* | [0.77] | [0.21] | [2.55]* | [0.19] |
| CHILD TO SCHOOL, | 1.53 | 0.0922 | 2.59 | 1.379 | 1.601 | -2.344 | 94.94 | 107.3 | -399.4 |
| SOME PRESCHOOLERS | (1.42) | (0.04) | (0.34) | (0.95) | (0.54) | (0.22) | (1.35) | (0.74) | (0.79) |
| | [1.42] | [0.03] | [0.32] | [1.01] | [0.37] | [0.13] | [1.42] | [0.55] | [0.56] |
| CHILD FINISHES | 1.078 | 3.721 | 14.56 | -0.4458 | 4.005 | -2.886 | 23.63 | 381.6 | 428.5 |
| SCHOOL, OTHER | (1.02) | (1.72) | (1.74) | (0.31) | (1.37) | (0.26) | (0.34) | (2.7)* | (0.79) |
| CHILDREN | [1.04] | [1.38] | [1.61] | [0.30] | [1.13] | [0.26] | [0.35] | [2.5]* | [0.84] |
| CHILD FINISHES | 4.566 | 12.75 | 37.94 | 5.017 | 2.319 | 44.30 | 298.9 | 473.0 | 2692.6 |
| SCHOOL, NO | (1.93) | (3.29)* | (1.92) | (1.57) | (0.44) | (1.66) | (1.94) | (1.87) | (2.08) |
| OTHER CHILDREN | [2.2] | [3.00]* | [2.51] | [1.35] | [0.31] | [1.52] | [1.86] | [1.33] | [2.18] |
| SPOUSE GETS RID | 0.0993 | 3.815 | -18.86 | -4.672 | 2.057 | -46.20 | -92.32 | 136.9 | -1994.1 |
| OF DISABILITY | (0.07) | (1.26) | (2.07) | (2.53) | (0.50) | (3.75) | (1.04) | (0.69) | (3.35) |
| THAT LIMITS WORK | [0.06] | [1.02] | [1.61] | [2.37] | [0.44] | [3.78] | [0.87] | [0.65] | [2.98] |
| SPOUSE ACQUIRES | -2.122 | 3.866 | -3.73 | -1.013 | -2.171 | 0.3802 | -101.8 | 53.32 | -198.5 |
| DISABILITY | (1.79) | (1.66)* | (0.48) | (0.63) | (0.69) | (0.04) | (1.31) | (0.35) | (0.39) |
| THAT LIMITS WORK | [2.39] | [1.35]* | [0.54] | [0.71] | [0.65] | [0.04] | [1.66] | [0.28] | [0.43] |
| Δ(OTHER | -0.0045 | 0.2871 | -0.2224 | 0.052 | 0.1741 | -0.6738 | 2.198 | 17.87 | -30.59 |
| INCOME)/1000 | (0.12) | (3.17)* | (0.91) | (0.99) | (1.42) | (2.05) | (0.87) | (3.02)* | (1.92) |
| | [0.09] | [2.64]* | [0.49] | [0.95] | [1.23] | [1.85] | [0.77] | [2.54]* | [1.42] |
| Δ(SPOUSE'S HOURS | 0.0093 | -0.1761 | 0.2337 | 0.054 | 0.0536 | 0.3028 | 3.11 | 0.2053 | 31.66 |
| UNEMPLOYMENT)/100 | (0.10) | (1.09) | (0.44) | (0.44) | (0.25) | (0.42) | (0.53) | (0.02) | (0.91) |
| | [0.15] | [0.86] | [0.46] | [0.36] | [0.22] | [0.37] | [0.51] | [0.02] | [0.91] |

Note: See Note to Table 2 for other variables included in equations and sample selection. The quit probability was also included. The quit probability was measured as the individual's deviation from the sample mean. The sample mean for P(Quit)=.1837. A "*" by a t-statistic indicates that the coefficient for quits is significantly different from the coefficient for non-quits, at the 5% level or better.

TABLE 4
HOURS CHANGE EQUATIONS
EFFECTS FOR NON-QUITTERS AND QUITTERS EVALUATED AT
SAMPLE MEAN FOR THE QUIT PROBABILITY

| | ΔHOURS/WEEK | | ΔWEEKS/YEAR | | ΔHOURS/YEAR | |
|---|---|---|---|---|---|---|
| | NonQuits | Quits | NonQuits | Quits | NonQuits | Quits |
| BIRTH, NO | -5.66 | -10.46 | -5.3140 | -6.894 | -455.27 | -576.76 |
| OTHER CHILDREN | (2.55) | (4.14) | (1.78) | (2.03) | (3.16) | (3.52) |
| | [2.82] | [2.82] | [1.92] | [2.23] | [3.27] | [3.00] |
| BIRTH, OTHER | 1.344 | 0.9648 | 1.601 | 1.344 | 113.12 | 128.72 |
| PRESCHOOLERS, NO | (0.76) | (0.36) | (0.68) | (0.37) | (0.99) | (0.73) |
| SCHOOL AGE CHILDREN | [0.98] | [0.35] | [0.77] | [0.35] | [1.41] | [0.75] |
| BIRTH, NO OTHER | 0.5213 | -2.597 | -4.168 | -5.033 | -102.95 | -242.75 |
| PRESCHOOLERS, SOME | (0.26) | (0.63) | (1.56) | (0.91) | (0.80) | (0.91) |
| SCHOOL AGE CHILDREN | [0.34] | [1.09] | [2.01] | [1.36] | [1.09] | [1.26] |
| BIRTH, OTHER | -2.981 | -12.43 | -2.872 | -13.09 | -188.33 | -623.63 |
| PRESCHOOLERS AND | (1.55) | (2.80) | (1.11) | (2.19) | (1.51) | (2.16) |
| SCHOOL AGE CHILDREN | [1.50] | [2.89] | [1.06] | [1.63] | [1.36] | [1.83] |
| CHILD TO SCHOOL, | -1.158 | 0.7914 | 2.712 | 10.46 | 37.69 | 347.21 |
| NO PRESCHOOLERS | (1.07) | (0.27) | (1.87) | (2.69) | (0.54) | (1.85) |
| | [1.20] | [0.35] | [1.59] | [2.1] | [0.48] | [1.57] |
| CHILD TO SCHOOL, | 1.190 | 1.108 | 0.668 | 3.626 | 60.03 | 183.69 |
| SOME PRE- | (1.10) | (0.46) | (0.46) | (1.13) | (0.86) | (1.18) |
| SCHOOLERS | [1.12] | [0.38] | [0.51] | [0.82] | [0.93] | [0.89] |
| CHILD FINISHES | 1.325 | 2.929 | 1.520 | 3.150 | 117.39 | 331.02 |
| SCHOOL, OTHER | (0.95) | (1.33) | (0.81) | (1.07) | (1.30) | (2.32) |
| CHILDREN | [0.97] | [1.05] | [0.80] | [0.88] | [1.44] | [2.22] |
| CHILD FINISHES | 4.665 | 7.775 | 5.163 | -6.082 | 293.38 | 121.80 |
| SCHOOL, NO | (1.92) | (1.18) | (1.58) | (0.69) | (1.86) | (0.28) |
| OTHER CHILDREN | [2.19] | [0.99] | [1.35] | [0.54] | [1.79] | [0.18] |
| SPOUSE GETS RID | -0.5019 | 1.242 | -5.651 | 0.3965 | -145.67 | 11.67 |
| OF DISABILITY | (0.36) | (0.37) | (3.04) | (0.09) | (1.62) | (0.05) |
| THAT LIMITS WORK | [0.29] | [0.32] | [2.65] | [0.08] | [1.28] | [0.05] |
| SPOUSE ACQUIRES | -2.663 | 3.158 | -2.296 | -5.777 | -164.82 | -109.79 |
| DISABILITY | (2.22) | (1.06) | (1.43) | (1.44) | (2.12) | (0.57) |
| THAT LIMITS WORK | [2.97] | [0.74] | [1.58] | [1.64] | [2.67] | [0.42] |
| Δ(OTHER | 0.0018 | 0.2795 | 0.0961 | 0.2572 | 4.32 | 21.13 |
| INCOME)/1000 | (0.04) | (2.86) | (1.76) | (1.96) | (1.64) | (3.33) |
| | [0.03] | [2.5] | [1.66] | [1.75] | [1.40] | [2.86] |
| Δ(SPOUSE'S HOURS | 0.0201 | -0.0964 | 0.1243 | 0.2149 | 5.835 | 4.536 |
| UNEMPLOYMENT)/100 | (0.22) | (0.54) | (1.02) | (0.89) | (0.99) | (0.39) |
| | [0.29] | [0.42] | [0.82] | [0.83] | [0.91] | [0.3] |

Notes: The results in this Table correspond to eq. 14 in the text. Interactions of all change variables with the quit probability, with quit, and with the quit probability times quit, were included. Also included were variables listed in the Note to Table 2. The numbers in the Table are the effects of a change in the exogenous variables on hours for quitters and non-quitters, evaluated at the sample mean of the quit probability.