LABOR SUPPLY WITH A MINIMUM HOURS THRESHOLD

DAVID CARD

Princeton University*

This paper considers the importance of minimum hours thresholds for the interpretation of individual labor supply data.   An analysis of quarterly labor-supply outcomes for prime-age males in the Survey of Income and Program Participation suggests that such thresholds are an important aspect of weekly hours choices.   A simple contracting model is presented that incorporates mobility costs and a nonconvexity in the relation between weekly hours and effective labor input.   This nonconvexity gives rise to a two-part employment schedule. In periods of low demand, some individuals are temporarily laid off, while others work a minimum threshold level of hours.   In periods of higher demand all available workers are employed at hours in excess of the threshold level.   The model provides a simple interpretation for the role of demand-side variables in explaining annual labor-supply outcomes.   It can also explain the weak correlations between annual hours and average hourly earnings that have emerged in earlier studies.   Under suitable assumptions on preferences, the intertemporal labor-supply elasticity can be recovered from the relationship between earnings and hours per week. Estimation results for the SIPP panel yield elasticity estimates that are similar to those in the literature based on annual data. If the model is correct, however, annual labor supply is considerably more sensitive to changes in productivity than these estimates suggest.

Following the lead of Hansen (1985) and Rogerson (1988), real business cycle theorists have recently recognized the importance of distinguishing between changes in the intensity of work effort per period and changes in the probability of employment.[1] Although the importance of the participation decision is widely acknowledged in studies of individual labor supply, most of the literature considers the question of participation at the

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[1]See the recent survey by Plosser (1989).

annual level.[2] There is surprisingly little research on the decomposition of annual hours into hours per week and weeks per year. Nevertheless, if most of the variation in annual hours consists of changes in weeks worked, with little or no change in average weekly hours, then one might not expect a systematic correlation between changes in annual hours and changes in average hourly earnings.

Unfortunately, most of the available longitudinal data sets are poorly suited to analyzing the components of intra-year variation in labor supply. Typically, these data sets are based on annual questionnaires that inquire about weeks of work and usual or average hours per week during the previous 12 months. The recently released Survey of Income and Program Participation (SIPP) is an important exception. Unlike other data sets, the SIPP questionnaire is administered every four months. The SIPP panel therefore provides a unique opportunity to analyze the composition of higher-frequency changes in labor supply.

This paper begins with a descriptive analysis of the components of variance of quarterly labor supply for adult men in the SIPP data set. Movements in and out of employment, changes in the number of weeks worked, and changes in the number of hours per week are all found to be important sources of variation in individual labor supply. Although hours per week vary, tabulations of the range of weekly hours suggest that most jobs are associated with a minimum-hours threshold. Over 90 percent of individuals observed on the same job during the sample period work 35 hours or more per week whenever they are employed.

With these facts in mind I present a simple contractual model of labor supply with a minimum-hours threshold. Following Rosen (1985), the model assumes a nonconvexity in the relation between hours per worker and effective labor input. This nonconvexity generates a two-part hours schedule. In periods of low employment demand some individuals are temporarily laid off, while others work a minimum-threshold level of hours. In periods of high demand all available workers are employed at hours in excess of the minimum threshold. The model provides a simple explanation for the finding of Ham (1986) that measures of annual labor supply vary with demand-side indicators, even controlling for wages. According to the model, changes in the probability of employment in any particular week vary with employer-

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[2]The issue of participation is widely analyzed in studies of female labor supply in particular. See the survey by Heckman and Killingsworth (1986).

demand conditions. Until all available workers are employed, however, aver-
age hourly earnings are constant. Thus, changes in weeks worked are corre-
lated with demand indicators, but may be independent of average hourly
earnings.

Under some restrictive (but standard) assumptions on preferences, the
weekly-earnings function implied by the model generates the same wage
elasticity of hours per week as a conventional life-cycle labor supply
model. Thus, the intertemporal labor-supply elasticity can be estimated by
considering the relation of average hourly earnings to hours per week,
although not necessarily to weeks worked.

Labor-supply equations for hours per week, weeks worked, and hours per
quarter are then estimated on the SIPP data. The estimation methods control
for measurement error in average hourly earnings and selection biases
generated by conditioning on employment status. Perhaps surprisingly,
however, I obtain estimates of the intertemporal substitution elasticity in
the same range as the previous literature.

I. PRELIMINARY DATA ANALYSIS

To investigate the characteristics of short-term fluctuations in
individual hours I have assembled a panel of observations for males age 22-
62 from the survey of Income and Program Participation (SIPP). The SIPP
has two important advantages over other longitudinal data sets. First, the
SIPP questionnaire is administered every four months.[3] Thus, unlike the
annual surveys in the Panel Study of Income Dynamics (PSID) and the
National Longitudinal Survey (NLS), the SIPP interview schedule permits a
detailed investigation of within-year variation in labor supply. Second,
the SIPP is much larger than other longitudinal surveys: 20,000 households
versus approximately 5000 in the PSID. These advantages are tempered by
the much shorter time dimension of the SIPP panel. Each household is
interviewed 8 or 9 times over a period of 36 months. Despite this
limitation, the size of the panel and frequency of observation make SIPP an
attractive data source for analyzing short-run hours determination.

The Census Bureau does not yet make available a complete longitudinal

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[3]See U.S. Bureau of the Census "Survey of Income and Program Participation Users' Guide"
(1987) for complete details on the SIPP sampling frame and interview schedules.

version of the SIPP data. I have therefore created my own longitudinal panel from the available cross-sectional samples.[4] The details of the merging procedure are described in the Data Appendix. The sample analyzed in this paper is drawn from the set of men with complete longitudinal information for all waves of the 1984 SIPP panel. To minimize problems associated with imputation errors, individuals with imputed responses for earnings, usual hours per week, or weeks worked in any month of the sample have been excluded.[5] A comparison of the characteristics of individuals with and without data imputations suggests that the biases arising from this exclusion are small.[6] Finally, I have restricted the sample to individuals who worked for pay at least one week between October 1983 and December 1985. These combined exclusions generate a usable sample of 4814 men.

Although weeks and earnings data are recorded on a monthly basis in the SIPP, hours per week are only recorded once for each employer over the previous four months. It is therefore impossible to recover independent observations on monthly labor supply, and I have aggregated the monthly information into quarterly data. An additional advantage of the quarterly time frame is that each calendar quarter has exactly 13 weeks, whereas calendar months may contain either 4 or 5 weeks. The staggering of interview dates for the four rotation groups in the SIPP panel implies that complete quarterly observations are only available from 1983-IV to 1985-IV. The final data set therefore contains 43,326 observations on the labor-supply outcomes of 4814 men in each of 9 quarters.

Table 1 presents information on the demographic characteristics and work histories of the sample. Data are presented for the sample as a whole and for the subsample of individuals who were employed in the first and last quarter of the sample period and who report no change in their

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[4]I am grateful to Kristin Butcher for her assistance with this task.

[5]Lillard, Smith and Welch (1986) discuss the limitations and biases in the Census Bureau "hotdeck" data imputation procedure. Their findings suggest that inclusion of observations with imputed earnings leads to significant measurement error.

[6]The most common imputation is for earnings. Roughly 15 percent of individuals in each survey decline to provide earnings information. Comparisons of the characteristics of individuals with and without imputations are presented in the Data Appendix.

Table 1

Characteristics of SIPP Sample

	Overall Sample	One-Employer Subsample
Demographic Characteristics		
1. Average Age (start of panel)	36.9	38.4
2. Percent Nonwhite	11.0	9.6
3. Average Education (start of panel)	13.1	13.1
Employment Data for Main Job		
4. Average Weeks Worked per Quarter	11.2	12.8
5. Average Hours per Week (in Quarters Worked)	41.1	42.6
6. Average Hours Worked per Quarter	472.1	546.1
7. Average Probability of Working in a Week (Over 117 Weeks)	85.8	98.2
8. Average Probability of Working in a Quarter (Over 9 Quarters)	90.1	99.2
9. Average Probability of Working in a Year (Over 2 Years)	96.1	99.9
10. Average Hourly Earnings 1984	10.40	11.10
11. Average Hourly Earnings 1985	10.65	11.32
12. Sample Size	4814	2864

Note: Sample consists of males age 22-59 as of Wave I of the 1984 SIPP Panel, with no imputations for earnings or hours in any wave and at least one week of paid work between October 1983 and December 1985. See Data Appendix. Individuals are followed for 9 quarters (1983-IV to 1985-IV).

employer during the 9 waves of the SIPP panel.[7] Approximately 60 percent of individuals fall into the latter category.

The demographic characteristics of the sample are very similar to those of a sample of similarly-aged men from the March Current Population Survey (CPS) who report working at some time in the previous year.[8] Rows 4-11 of the table contain information gleaned from earnings and hours information reported for each individual's main job. Individuals report working an average of 11.2 weeks per quarter, and just over 41 hours per week in those weeks in which they work. Average hourly earnings, reported in rows 10 and 11 of the table, are again similar to those of men in the same age group in the CPS.[9]

The average employment rate varies inversely with the length of the interval used to define employment status. Over the 27-month sample period 85.8 percent of the sample were employed in any given week, 90.1 percent in any given quarter, and 96.1 percent in either 1983 or 1984.[10] Movements in and out of employment are clearly a more important source of variation in individual hours at higher frequencies of observation.

The contribution of changes in employment status to the total variation in quarterly hours of work is addressed in Table 2. The total variance in quarterly hours around individual-specific means is presented in row 1. This variance is defined as

$$v = \frac{1}{N \cdot T} \Sigma_i \Sigma_t \ (H_{it} - H_i)^2,$$

where H_{it} represents the hours worked by individual i in period t, H_i represents the mean hours worked per quarter by i, N represents the number of individuals in the sample, and T (≈ 9) represents the number of time

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[7]Inspection of the sequences of employer identifiers reported by individuals suggests that there may be both false transitions and unreported changes in the data. Census Bureau researchers report some difficulty with the coding of employer identities following a spell of unemployment (U.S. Bureau of the Census (1989), pp. 16-18).

[8]For example, among men age 22-60 in the March 1985 CPS who worked in the previous year, the average age is 37.2 years, the average years of completed education is 13.0, and the percent nonwhite is 11.2.

[9]Average hourly earnings in the March 1985 CPS sample of men age 22-60 who worked in 1984 are $10.76.

[10]Recall that individuals had to work some time between October 1983 and December 1985 in order to appear in the sample.

Table 2

Components of Variance of Quarterly Hours Around
Individual-Specific Means
(Standard Errors in Parentheses)

		Overall Sample	One-Employer Subsample
1.	Unconditional Variance of Hours	16586.8 (369.9)	5766.9 (257.3)
2.	Conditional Variance of Hours	10655.5 (254.1)	4371.6 (194.7)
3.	Percent of Unconditional Variance Due to Probability of Employment	48.7	27.0
4.	Conditional Variance of Log Hours (x100)	14.67 (.45)	3.90 (.27)
5.	Conditional Variance of Log Hours/Week (x100)	4.06 (.15)	1.45 (.10)
6.	Conditional Variance of Log Weeks (x100)	6.44 (.21)	1.33 (.11)
7.	Conditional Covariance of Log Hours/Week and Log Weeks (x100)	2.08 (.10)	.56 (.06)

Note: See text. All data are deviated from individual-specific means.

periods per individual.[11] In the overall sample $v = 15,587$, while within the one-employer subsample $v = 5756$. These estimates imply standard deviations of 129 hours and 76 hours, respectively. The corresponding coefficients of variation are .27 and .14.

For a given individual, the variance of hours can be decomposed into a component due to the conditional variance of hours, given employment, and a component due to changes in employment status. Specifically, let p_i represent the probability that i is employed in any quarter, let H_i^C represent the conditional mean of hours, given that i is employed, and let v_i^C represent the conditional variance of i's hours, given that i is employed. Finally, let v_i represent the unconditional variance of i's quarterly hours. Then

$$v_i = p_i \, v_i^C + p_i \, (1 - p_i) \, (H_i^C)^2 .$$

The first term can be interpreted as the part of v_i due to the conditional variance of hours, given employment, while the second component can be interpreted as the part of v_i due to the variance of employment outcomes. For the entire sample:

$$V = \frac{1}{N} \, \Sigma_i \, v_i = \frac{1}{N} \, \Sigma_i \, \{ p_i \, v_i^C + p_i \, (1 - p_i) \, (H_i^C)^{\,2} \},$$

$$= \frac{1}{N} \, \Sigma_i \, p_i \, v_i^C + \frac{1}{N} \, \Sigma_i \, p_i (1 - p_i)(H_i^C)^{\,2} .$$

The share of the average variance in hours due to the variability of employment outcomes is then

$$s \equiv \frac{\frac{1}{N} \, \Sigma_i \, p_i (1 - p_i)(H_i^C)^2}{v} = \frac{1}{N} \, \Sigma_i \, \alpha_i \, s_i ,$$

where $s_i = p_i \, (1-p_i)(H_i^C)^2 \, / \, v_i$ is the share of v_i attributable to changes over time in i's employment status and $\alpha_i = v_i/v$ is the relative variability of hours for individual i. The value of this share in the

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[11]The degrees of freedom of the estimated sample variance are not adjusted for the estimated individual means. This adjustment would reduce the estimated variances in Table 2 by 11 percent.

overall sample is just under 50 percent. Thus, quarter-to-quarter movements in and out of employment contribute about as much to the average variability of hours as changes in the level of hours, conditional on employment.[12] In the one-employer subsample the share is lower, as could be expected given the much higher employment probabilities for this group.

The conditional variance of hours per quarter can itself be decomposed into a component due to the variation in weeks per quarter, a component due to the variation in hours per week, and a covariance term. One such decomposition is presented in Table 2, based on a logarithmic transformation of conditional hours. Let n_{it} represent the number of weeks of employment for individual i in quarter t, let h_{it} represent hours per week reported by i in quarter t, and observe that $H_{it} = n_{it} h_{it}$. The conditional variance of log hours for individual i can be written as:

$$var (\log H_{it}|H_{it} > 0) = var (\log n_{it}|H_{it} > 0) + var (\log n_{it}|H_{it} > 0)$$

$$+ 2cov (\log n_{it}, \log h_{it} \mid H_{it} > 0).$$

Sample averages of these 4 terms are presented in rows 7-11 of Table 2.[13] In both the overall sample and the one-employer subsample the covariance of weeks worked and hours per week is positive. If the covariance is attributed equally to weeks and hours per week, the share of the conditional variance of hours due to changes in weeks worked within a quarter is 60 percent in the overall sample and 50 percent in the subsample.

The data in Table 2 suggest that changes in quarterly labor supply for a given individual arise from three sources: movements in and out of employment for an entire quarter; changes in the numbers of weeks worked, conditional on employment in the quarter; and changes in the number of hours per week. For individuals who worked at the same job during the sample period, the contributions of these three components are about equal. For others, the largest share of the unconditional variance is attributable to movements in and out of employment, and the smallest share to changes in

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[12]This share is roughly similar to the share of the variance of aggregate quarterly hours attributable to changes in employment: see Hansen (1985, p. 311) and Coleman (1984).

[13]Again the data on weeks and hours per week are adjusted for individual-specific means.

hours per week.[14]

Further insight into the nature of hours variation is provided by the cross-tabulations in Table 3. Each row of the table gives the percentage of individuals with a particular range of hours per week during the sample period. The first 4 rows give the percentages of individuals who report the same hours per week in all 9 quarters, while the next rows give the percentages of individuals with some variation in weekly hours. In the overall sample 30 percent of individuals report constant weekly hours: the vast majority of these report exactly 40 hours per week. Among individuals with some variation in hours per week, two cases are prominent. About 40 percent report a minimum of exactly 40 hours. Another 40 percent report a range of hours with a minimum below 35 hours.

The distribution of the range of weekly hours is slightly different in the one-employer subsample. An interesting aspect of the table, however, is the fact that the distributions in the overall sample and the one-employer subsample are very similar, conditional on the presence or absence of weeks of nonemployment. Thus the distribution in column 4 is very similar to that in column 1, while the distribution in column 5 is similar to that in column 2. The one-employer group differs from the overall sample mainly in the likelihood of lost weeks.

These tabulations suggest that some notion of a minimum threshold in the weekly hours of prime-age males is probably useful. Three-quarters of the overall sample and almost 90 percent of the one-employer subsample work at least 35 hours per week whenever they are employed. Despite these tendencies, a significant fraction of individuals fall below 35 hours per week at some point during the sample period. Low weekly hours are particularly likely among individuals who suffer weeks of nonemployment: one-half of these fall below 35 hours at some point. Thus although a rigid weekly-hours threshold is overly restrictive, it is a useful first approximation, at least within jobs.

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[14]These calculations should be interpreted carefully, since differential measurement errors in weeks worked and hours per week can lead to over- or understatement of the relative contributions.

Table 3

Individual-Specific Range in Hours Per Week

	Overall Sample			One-Employer Subsample		
	No Weeks Lost	Some Weeks Lost	All	No Weeks Lost	Some Weeks Lost	All
	(1)	(2)	(3)	(4)	(5)	(6)
Percent with:						
1. Constant Hours Per Week						
(a) 40 hours per week	35.3	14.5	26.3	36.5	20.2	34.4
(b) 35-39 hours per week	1.2	.5	.9	1.3	.6	1.2
(c) <35 hours per week	.2	3.4	1.6	.2	0	.1
(d) >40 hours per week	1.4	.9	1.1	1.5	0	1.3
2. Variable Hours Per Week						
(a) Minimum 40 hours per week	34.2	15.7	26.2	33.8	16.0	31.6
(b) Minimum 35-39 hours per week	7.8	6.3	7.2	7.6	5.2	7.3
(c) Minimum <35 hours per week	5.8	53.6	26.6	5.2	54.4	11.4
(d) Minimum >40 hours per week	14.2	5.2	10.3	14.0	3.6	12.6
3. Number of Individuals	2719	2095	4814	2502	362	2864

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II. A SIMPLE CONTRACTING MODEL WITH MINIMUM HOURS

This section presents a simple contracting model of intertemporal labor supply that captures some of the features highlighted in the previous section, while at the same time addressing an important issue raised by Ham's (1986) analysis of intertemporal labor supply. Ham's findings, and the results presented below, suggest that industry-specific employment-demand variables are important determinants of individual labor supply, even controlling for wages. Ham interprets this finding as evidence against the strict life-cycle labor-supply model.[15] In the model presented here this finding is interpreted as evidence of a nonconvexity in the relation between individual hours and output. As noted by Rosen (1985), this nonconvexity gives rise to a two-part employment contract. In low-demand states, some employees work "standard hours," while others are temporarily laid off. In high-demand states, all available employees work "overtime hours," at a premium that depends on the intertemporal elasticity of labor supply. Such a two-part contract is useful to explain the prominent spike in the distribution of weekly hours at 35-40 hours. It also gives a straightforward interpretation of demand-side variables in the labor-supply equation. These appear as determinants of the probability of employment in any particular week. Finally, to the extent that changes in annual hours take place through changes in the number of weeks of employment at standard hours, the two-part employment contract explains the rather weak correlations between annual hours and average hourly earnings that have emerged in previous studies.[16]

The model is a straightforward re-interpretation of the model presented by Rosen (1985, Section V). Consider a multi-period economy with firm-specific demand or productivity shocks in which workers attach themselves permanently to firms. In each period a firm selects the number of employed workers, n_t, and hours per employed worker, h_t. Revenues of a particular firm are

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[15]See Card (1987) and Heckman and MaCurdy (1988) for further discussion of these findings and some of the objections that have been raised to their interpretation.

[16]See Altonji (1986), for example. Hansen (1985) and Rogerson (1988) point out that nonconvexities can explain the co-existence of arbitrarily elastic intertemporal labor supply with relatively weak correlations between hours of work and average hourly earnings.

$$R(n_t f(h_t), \theta_t)$$

where $n_t f(h_t)$ represents effective labor input in t, and θ_t represents a demand or productivity shock. The correct time period in this context is the minimum period over which employment is fixed, and corresponds most naturally to a week. Assume that f exhibits first increasing and then decreasing returns to hours. A convenient example of such a function is one with "start-up" costs:

$$f(h_t) = 0, \qquad\qquad h_t < h_0,$$

$$= (h_t - h_0)^\delta, \qquad h_t \geq h_0, \quad 0 < \delta < 1.$$

Start-up costs, or similar nonconvexities in the relation between hours per worker and effective labor input, give rise to a minimum-hours threshold (see below). Fixed costs of travelling to work or nonconvexities in tastes for leisure could generate the same phenomenon. I suspect, however, that an important component of threshold-like behavior is attributable to technology.

Assume that workers' preferences between hours of work and consumption are additively separable over time. In addition, assume that the within-period utility function U is additively separable in consumption, c_t, and hours of work:

$$U(c_t, h_t) = u(c_t) - \phi(h_t),$$

where u and ϕ are strictly increasing, u is concave, ϕ is convex, and $\phi(0)=0$. Assume that a pool of workers of size n_0 is initially attached to the firm. In each period, a random selection n_t of these are actually employed ($n_t \leq n_0$), and the remainder are temporarily laid off. The expected utility earned by a representative worker is therefore

$$n_t/n_0 \ U(c_t, h_t) + (1 - n_t/n_0) \ U(\bar{c}_t, 0),$$

where c_t denotes consumption in period t if employed, h_t represents hours per employed worker in t, and $\bar{c}_t$ represents the consumption of unemployed workers.

Suppose that training and recruiting costs per worker are τ, and that workers and the firm share a common discount rate β. An optimal contract

is a set of contingent functions $c_t(\theta_t)$, $\bar{c}_t(\theta_t)$, $n_t(\theta_t)$, and $h_t(\theta_t)$, and an initial firm size n_0 that maximizes

$$E \; \Sigma_t \; \beta^t \; \{R(n_t f(h_t), \; \theta_t) - n_t c_t - (n_0 - n_t)\bar{c}_t\} - \tau n_0,$$

subject to

$$E \; \Sigma_t \; \beta^t \; \{n_t U(c_t, \; h_t) + (n_0 - n_t)U(\bar{c}_t, \; 0)\} \geq n_0 V^*,$$

and

$$n_t \leq n_0 \text{ for all t.}$$

Here, E denotes expectations over the joint distribution of $\{\theta_t, \; t=0,\ldots\infty\}$, taken with respect to information available in an initial recruiting period, and V^* represents the equilibrium value of alternative contracts available to each worker.

Let ψ represent the per-worker value of the multiplier associated with the expected utility constraint, and let $\beta^t \mu(\theta_t)$ represent the present value of the multiplier associated with the maximum employment constraint in period t when the demand/productivity shock is θ_t. The first-order conditions for an optimal contract imply that consumption is constant over time and states and independent of employment or unemployment status:

$$c_t(\theta_t) = \bar{c}_t(\theta_t) = [u']^{-1} \; (\psi^{-1}).$$

The first-order conditions for employment and hours are

$$f(h_t) \; R_1(n_t f(h_t), \; \theta_t) = \psi \cdot \phi(h_t) + \mu_t(\theta_t), \tag{1}$$

$$f'(h_t) \; R_1(n_t f(h_t), \; \theta_t) = \psi \cdot \phi'(h_t). \tag{2}$$

In periods when $n_t < n_0$, $\mu_t = 0$ and these equations can be rewritten as

$$\frac{f'(h_t)}{f(h_t)} = \frac{\phi'(h_t)}{\phi(h_t)} \, ,$$

implying that $h_t = h^*$, independent of θ_t.[17] If the marginal revenue function is monotonic in θ, then there exists a critical level of the demand/productivity shock, say θ^*, such that if $\theta_t < \theta^*$ then $h_t = h^*$ and $n_t < n_0$, while if $\theta_t > \theta^*$ then $n_t = n_0$ and $h_t > h^*$. In low-demand states, hours per worker are fixed and marginal adjustments to the effective labor force are met by changing the level of employment. In high-demand states, on the other hand, employment is constrained by the size of the available pool, and marginal changes in the effective labor force are accomplished by varying hours per worker.

An important aspect of this contract is its decentralizability. On the demand side, if the firm acts unilaterally to maximize profits, subject to an earnings schedule $g(h_t)$ for each employed worker, then the firm will make the jointly optimal employment and hours choices provided that $g(h_t) = \psi \cdot \phi(h_t)$, where ψ is the multiplier from the associated joint optimization.[18] On the supply side, if workers have access to risk-neutral savings/insurance markets, and face a contractual earnings function $g(h_t) = \psi \cdot \phi(h_t)$, then there exists a once-for-all transfer payment from the firm that causes workers to supply the jointly optimal hours choice h_t if they are offered employment in period t, and to unilaterally implement the jointly optimal consumption choices.[19] To see this, consider the problem of choosing consumption if employed c_t, hours if offered employment h_t, and consumption if unemployed $\tilde{c}_t$ to maximize

$$E \, \Sigma_t \, \beta^t \, \{ n_t/n_0 \, U(c_t, \, h_t) + (1 - n_t/n_0) \, U(\tilde{c}_t, \, 0) \}$$

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[17]The existence of an hours choice h^* satisfying (1) and (2) when $\mu_t = 0$ requires that f exhibit first increasing and then decreasing returns to scale. For example, if $f(h)=(h-h_0)^\delta$ and $\phi(h)=1/\epsilon \, h^\epsilon$, with $\epsilon>1>\delta$, then $h^* = \epsilon/(\epsilon-\delta) \, h_0$.

[18]This is readily seen by considering the maximization of the discounted present value of revenues less wage costs, subject to the earnings function $g(h_t) = \psi \cdot \phi(h_t)$ and the constraint $n_t \le n_0$.

[19]In states of less than full employment, individual workers must still rely on the firm (or some randomization device) to choose which workers are employed and which are not.

subject to

$$E \sum_t \beta^t \{n_t/n_0 \, c_t + (1 - n_t/n_0) \, \tilde{c}_t - n_t/n_0 \, g(h_t)\} = B,$$

where B is an initial transfer payment from the firm to each worker, and expectations are taken with respect to the distribution of employment levels induced by the jointly optimal contract.[20]

Let λ denote the multiplier associated with the lifetime wealth constraint. The first-order conditions for this problem include

$$u'(c_t) = u'(\tilde{c}_t) = \lambda$$

and

$$\phi'(h_t) = \lambda \, g'(h_t).$$

If B is selected to allow the same consumption stream as in the jointly optimal contract, then $\lambda = \psi^{-1}$, the inverse of the multiplier associated with the per-worker utility constraint in the optimal contract. Assuming that the contractual earnings function is $g(h_t) = \psi \cdot \phi(h_t)$, the first-order condition for hours is satisfied trivially for any h_t.[21] Thus the jointly optimal contract is supported by a simple earnings function that is proportional to the disutility of hours function $\phi(h)$.

It is interesting to compare this contractual-earnings function to the earnings function generated by a conventional life-cycle labor supply problem, under a similar specification of preferences. In the usual life-cycle problem, earnings represent the product of hours of work and a parametric wage rate w_t. The first-order condition for the optimal choice of hours in period t is

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[20]This formulation ignores unemployment benefits or other income sources in the unemployed states. Unemployment benefits can be added with no change in the implications of the model.

[21]The contract leaves individuals who are employed indifferent between alternative hours choices in the sense that any level of hours will satisfy the first-order conditions. Unemployed individuals are in one sense "better off," because they have the same consumption level but more leisure. However, employed individuals are earning income, whose shadow value is λ per unit. Since $\lambda g(h_t) = \phi(h_t)$, the shadow value of earnings just offsets the value of foregone leisure.

$$\phi'(h_t) = \lambda_\ell w_t,$$

where λ_ℓ is the marginal utility of wealth in the life-cycle allocation problem. If λ_ℓ is set to give the same consumption stream as the jointly optimal contract, then $\phi_\ell = \psi^{-1}$. Under this assumption the conventional life-cycle earnings function $g_\ell(h_t)$ satisfies

$$g_\ell(h_t) = \psi\, h_t\, \phi'(h_t) > \psi\, \phi(h_t) = g(h_t),$$

since $\phi(h)$ is convex. However, if ϕ exhibits a constant elasticity, implying that the intertemporal substitution elasticity is constant, then the contractual-earnings function is proportional to the earnings function implied by the usual life-cycle labor-supply model.

Specifically, suppose that $\phi(h) = 1/\varepsilon\, h^\varepsilon$, where $\varepsilon = (1+\eta)/\eta$, and $\eta>0$. Then the contractual earnings function satisfies

$$\log h_t = \eta \log (g_t/h_t) + \text{constant},$$

where η is the conventional intertemporal substitution elasticity. Unlike the usual life-cycle model, however, this equation _only_ holds for hours worked in excess of standard hours h^*. If the firm's decision period is a week, then changes in weekly hours over and above standard hours are related to weekly average hourly earnings by a conventional life-cycle labor-supply equation. Monthly or annual hours bear no such simple relation to average hourly earnings. For example, if changes in annual hours consist of changes in weeks worked at standard hours, then hours will appear to vary at fixed wage rates. Research by Abowd and Card (1989) suggests that much of the measured variation in annual hours for prime-age males actually occurs at constant wages.

This model can also explain the role of demand-side variables (like employment growth rates for an individual's industry or region) in an annual labor-supply equation. According to the model, variation in weeks worked depends on the realization of demand shocks and is only indirectly related to average hourly earnings while employed. Demand-side variables therefore appear in an annual labor-supply equation as proxies for the determinants of the number of weeks worked. With mobility costs and a non-

convex production technology, it is impossible to specify the labor-supply equation independently of demand-side information.[22]

To summarize, this section presents a simple long-term contracting model with a nonconvexity in the relation between output and hours per worker. This feature gives rise to a two-part contractual-hours function. In periods of low employment demand, some individuals are temporarily laid off, while the remainder work standard hours. In these states, individuals shift between employment and unemployment at constant average hours earnings. In periods of high employment demand, all individuals are employed at hours greater than standard hours, with a positive relation between earnings and hours worked. Under a set of restrictive assumptions on preferences (including additive separability between consumption and leisure within periods), the contractual-earnings function generates the same elasticity between hours and average hourly earnings as a conventional life-cycle labor-supply model. This relation only holds for hours in excess of standard hours and need not hold for aggregates of hours (such as annual or quarterly totals).

It is worth noting that the model is highly restrictive. First, preferences are assumed to be additively separable within and across periods, with no allowance for heterogeneity across people or over time. Second, jobs are assumed to last indefinitely. Third, the form of the contractual-earnings schedule ignores institutional restrictions such as compulsory overtime legislation. I hope to be able to address some of these issues in future work.

III. MODEL ESTIMATION

This section applies the model of the previous section to quarterly changes in individual labor supply for prime-age males in the SIPP. According to the model, changes in hours between periods in which firm-level employment is fixed should be related to changes in average hourly earnings by a conventional intertemporal substitution elasticity. I interpret this period of fixity as a week. Changes in labor supply between longer periods (quarters or years) consist of changes in hours per week,

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[22]A similar argument shows how reported weeks of unemployment may be negatively correlated with reported annual hours.

which are necessarily correlated with average hourly earnings, and changes in weeks worked, which are not. Therefore, if the model is correct, one should expect larger elasticities between average hourly earnings and hours per week than between average hourly earnings and weeks worked.

Before proceeding to analyze quarterly data from the SIPP panel, it is worth checking whether <u>annual</u> labor-supply data from the SIPP sample show similar patterns to data from the PSID and NLS. Since much of the previous analysis of intertemporal labor supply has been conducted on first-differenced data, I aggregated quarterly information for the SIPP sample into annual data for 1983 and 1984, and computed the variances and covariances of changes in annual hours, annual earnings, and average hourly earnings.

In the SIPP sample, changes in the logarithm of annual hours are strongly negatively correlated with changes in the logarithm of average hourly earnings: the regression coefficient of the change in log hours on the change in log average hourly earnings is -.41, with a standard error of .03. When potential labor market experience is used as an instrumental variable for the change in wages, the estimated elasticity of annual hours with respect to average hourly earnings is .87, with a standard error of .38. These elasticities are not much different from those reported by Abowd and Card (1989) for samples of men drawn from the PSID and NLS.

Estimation of the simple model of the previous section is complicated by three issues -- measurement errors, sample selection bias, and aggregation bias. The importance of measurement error in the covariance properties of hours and average hourly earnings is well-documented.[23] Indeed, the strong negative correlations between changes in hours and changes in average hourly earnings that appear in virtually every panel data study are generally attributed to measurement error. There is less agreement on how to handle the problem. MaCurdy (1981) proposed the use of polynomials of age and education as instrumental variables for the change in average hourly earnings. Provided that tastes for leisure are uncorrelated with age and education, these are legitimate instruments. Even under this assumption, the explanatory power of such variables is extremely weak, often leading to imprecise estimates of the elasticity

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[23]See Altonji (1986), for example.

between hours and average hourly earnings.[24]

The use of quarterly observations provides an opportunity for an alternative instrument. Suppose that individual employment situations are characterized by a match-specific seasonal productivity component. Then changes in average hourly earnings will be positively correlated with changes that occurred 4 quarters ago (or that occur 4 quarters in the future). Among men in the one-employer subsample this individual-specific seasonal correlation is small, but highly statistically significant. A regression of the change in log average hourly earnings on the change 4 quarters in the past and a set of unrestricted quarterly dummy variables yields a coefficient of .035, with a standard error of .007.[25] An important limitation of this instrument is that it is unavailable for individuals who did not work 4 and 5 quarters in the past. In the one-employer subsample the fraction of quarterly observations affected by this problem is small (1.4 percent).

Compared to the issue of measurement error, the issue of selection bias has received relatively little attention in the literature on male labor supply.[26] With quarterly data, however, selection bias is a potentially serious problem, since many more individuals spend at least a full quarter unemployed than spend a whole year unemployed. To see the likely nature of the biases, consider the following equation for log hours per week of individual i in quarter t:

$$\log h_{it} = \alpha_i + \eta \log w_{it} + \nu_{it}. \tag{3}$$

Here, α_i refers to a person- and job-specific constant, w_{it} refers to average hourly earnings in the quarter, η is the intertemporal substitution elasticity, and ν_{it} is a stochastic disturbance incorporating measurement

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[24]This is evidenced by the rather large standard error associated with the instrumental variables estimates of the elasticity of changes in annual hours with respect to changes in average hourly earnings.

[25]Note that this seasonal correlation does not simply reflect seasonal tastes for leisure, since unrestricted quarterly dummies are included.

[26]Sample selection is routinely addressed in the literature on female labor supply. See Heckman and Killingsworth (1986) for a recent survey.

error. According to the model in Section II, the probability that i is employed in t, and therefore observed in the sample, depends only on the state of demand at i's employer. If i is employed, hours and earnings are related by a deterministic earnings function. On this strict interpretation v_{it} consists solely of measurement error, and there is no selection bias (i.e., $E(v_{it}|H_{it}>0)$ does not depend on w_{it}). If hours per week and the probability of employment share any discretionary components, however, one would expect v_{it} to be positively correlated with the likelihood that i is employed in t.[27] In this case, the sample-selection process leads to a negative bias in estimates of η.

Suppose that i is employed in period t if a latent random variable y_{it} is positive, where

$$y_{it} = \gamma_i + z_{it}\pi - \xi_{it}. \tag{4}$$

Here γ_i has the interpretation of a person- and job-specific fixed effect, z_{it} consists of measured covariates, and ξ_{it} is a transitory component. According to the model developed above, z_{it} should contain proxies for the strength of employment demand in i's industry in period t. In the application below, z_{it} consists of the level of employment in quarter t in i's 1-digit industry.

To the best of my knowledge there is no completely satisfactory method for handling the selection biases implied by (4) in the estimation of (3). The following strategy is based on a suggestion of Olsen (1980). Suppose that ξ_{it} is uniformly distributed on the interval [0,1]. Then p_{it}, the probability that i is employed in t, is given by

$$p_{it} = \gamma_i + z_{it}\pi.$$

The parameter(s) π can be estimated by a first-differenced linear-probability model applied to the actual sequence of employment indicators

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[27] For example, sickness or other sources of taste variation ignored in the model may lead individuals to reduce their hours when employed and to reduce the probability of employment.

for individual i. Suppose in addition that

$$E(\nu_{it}|\xi_{it}) = c_0 + c_1 \, \xi_{it},$$

i.e., that the conditional expectation of the error component ν_{it} is linear in ξ_{it}. Then

$$
\begin{aligned}
E(\log h_{it}|H_{it}>0) &= \alpha_i + \eta \log w_{it} + c_0 + c_1 \, E(\xi_{it}|\xi_{it} \le \gamma_i + z_{it}\pi) \\[2mm]
&= \alpha_i + \eta \log w_{it} + c_0 + c_1 \cdot (\gamma_i + z_{it}\pi)/2 \\[2mm]
&= (c_0 + \alpha_i + c_2 \cdot \gamma_i) + \eta \log w_{it} + c_2 \, (z_{it}\pi),
\end{aligned}
$$

where $c_2 = c_1/2$. Given an estimate of π, say $\bar{\pi}$, the parameter η can be consistently estimated by taking first-differences of observed hours choices (conditional on positive hours in t and $t-1$):

$$\Delta\log h_{it} = \eta \, \Delta\log w_{it} + c_2 \, (\Delta z_{it}\bar{\pi}) + e_{it}. \tag{5}$$

The assumptions underlying this procedure are quite restrictive. At the very least, however, estimates of c_2 provide simple evidence on the importance of sample-selection biases in the estimation of quarterly hours equations.

The final issue in estimation of the model is the fact that observations on the critical "hours per week" variable are only available quarterly. If hours per week are constant within a quarter, then the model applies directly and there is no aggregation bias. If hours per week vary within a quarter, however, two problems arise. First, recall biases and the wording of the SIPP question on hours per week make it conceivable that some individuals report an answer like "40 hours per week" rather than a

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precise average of weekly hours.[28]   Second, even if individuals correctly report an average of weekly hours, the convexity of the weekly earnings function implies that averages of hours per week are not necessarily related to average hourly earnings by (3).[29]   I suspect that the first of these problems is more important than the second.   Whatever the magnitude of reporting biases in the SIPP, however, they are presumably less troublesome than the biases in an annual survey.

Estimation results for various first-differenced labor-supply equations fit to observations for the one-employer subsample of the SIPP data set are presented in Table 4.   Each column of the table corresponds to a particular choice of estimation method, sample, and instrumental variable for the change in average hourly earnings.   For each choice, estimation results are presented for three measures of labor supply: the logarithm of total quarterly hours, the logarithm of hours per week, and the logarithm of weeks worked.   Because the logarithm of quarterly hours is the sum of the logarithms of hours per week and weeks worked, the coefficients in rows 4(i) and 5(i) sum to the coefficients in row 3(i).

Two samples are constructed from the quarterly labor-supply data of the one-employer SIPP sample: the set of all available quarterly changes in labor supply (denoted by the column heading "All"); and the subsample of changes for which a corresponding change 4 quarters in the past or 4 quarters in the future is available (denoted by the column heading "Subsample").   Use of $\Delta \log w_{it-4}$ or $\Delta \log w_{it+4}$ as an instrumental variable for $\Delta \log w_{it}$ is restricted to the subsample.

Ordinary least squares (OLS) estimation results are presented in columns (1) and (2).   The estimates are very similar in the two samples, and suggest that changes in hours per week and changes in weeks worked are both strongly negatively correlated with changes in average hourly earnings.   The estimated elasticity of total quarterly hours with respect to average hourly earnings is -.39, very similar to estimates in the literature based on annual data.   It is interesting to note that the

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[28]In fact, the SIPP interviewer manual (U.S. Bureau of the Census (1989)) instructs individuals with varying hours per week to report their "most frequent" weekly hours rather than an average of hours per week over the preceding months.

[29]This problem arises quite generally.   If the correct "period" for analyzing labor supply is shorter than the periodicity of the available data, the estimated labor-supply parameters will differ from the true parameters.

Table 4

Estimated Hours Equations: First Differenced Specifications
Fit to One-Employer Subsample
(Standard Errors in Parentheses)

| | Estimation Method and Sample | | | | | | | |
| | OLS | | IV | | | | | |
Sample:	All (1)	Sub-sample (2)	All (3)	All (4)	Sub-sample (5)	Sub-sample (6)	Sub-sample (7)	Sub-sample (8)
1. Instrumental Variable for Δw_t	--	--	Exp	Exp	Exp	Exp	Δw_{t-4}, Δw_{t+4}	Δw_{t-4}, Δw_{t+4}
2. Selection Correction	No	No	No	Yes	No	Yes	No	Yes
3. Log Hours Equation								
(i) Wage Coefficient	-.39 (.01)	-.38 (.01)	.00 (.50)	-.01 (.50)	-.01 (.52)	-.02 (.52)	.16 (.22)	.14 (.22)
(ii) Selection Coefficient	--	--	--	7.26 (1.45)	--	6.37 (1.58)	--	6.02 (1.27)
4. Log Hours/Week Education								
(i) Wage Coefficient	-.29 (.01)	-.28 (.01)	.18 (.36)	.18 (.36)	.22 (.38)	.21 (.38)	.10 (.14)	.10 (.14)
(ii) Selection Coefficient	--	--	--	2.78 (1.02)	--	2.55 (1.16)	--	2.80 (.79)
5. Log Weeks Equation								
(i) Wage Coefficient	-.10 (.01)	-.09 (.01)	-.18 (.32)	-.19 (.32)	-.23 (.33)	-.24 (.33)	.05 (.13)	.05 (.13)
(ii) Selection Coefficient	--	--	--	4.48 (.91)	--	3.83 (1.02)	--	3.21 (.76)

Notes: All sample contains 19842 observations. Subsample contains 19566 observations for which change in average hourly earnings four quarters in past/future is available. Selection correction is based on estimated first-differenced linear probability model: see text. In row 1, Exp refers to potential labor-market experience, Δw_t refers to the change in log average hourly earnings, and $\Delta w_{t-4}/\Delta w_{t+4}$ refer to changes in log average hourly earnings four quarters in past/future.

elasticity of hours per week is more negative than the elasticity of weeks. One simple explanation for this finding is that measurement errors are proportionally smaller in the weeks measure than in the measure of hours per week.[30] Given the nature of the SIPP questionnaire, this seems plausible.

Instrumental variables (IV) estimates that potentially eliminate the biases caused by measurement errors in hours are presented in columns (3) through (8). The estimates in columns (3)-(6) use potential labor-market experience (age minus education minus 5) as an instrument for the change in average hourly earnings. The use of this instrumental variable leads to similar results on the total sample and subsample. In both cases the IV estimate of the elasticity of hours per week with respect to average hourly earnings is positive but relatively imprecise. The IV estimates of the elasticity of weeks with respect to average hourly earnings are negative, but again relatively imprecise. The elasticities of hours per week and weeks tend to offset each other, leading to a wage elasticity of total quarterly hours that is close to zero.

Columns (4) and (6) combine IV estimation with the linear selection correction proposed in equation (3). The first-stage equation uses changes in the log of aggregate employment in the individual's one-digit industry as well as an unrestricted set of quarterly dummy variables to predict the change in the probability of employment. Industry-level employment changes are reasonably strong predictors of the change in the probability of employment: the t-statistic associated with the measured change in industry employment is 2.7. Again, estimation results are similar in the complete sample and the subsample: in all cases the estimated coefficient associated with the selection term is positive and statistically significant, although the addition of the selection term has little impact on the estimated wage elasticities. As expected, the selection term has a larger coefficient in the equation for the number of weeks worked than the number of hours per week, although one can reject the hypothesis that the number of hours per week is independent of industry demand, controlling for the level of average hourly earnings. The large and statistically significant estimates

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[30]Any measurement error in log hours or weeks induces a perfectly negatively correlated error in log average hourly earnings. The greater the variance in such measurement errors, the more negative is the OLS estimate of the elasticity between the hours measure and average hourly earnings.

of the selection coefficients confirm the finding of Ham (1986) that changes in demand-side variables influence individual labor supply, even controlling for individual wages.

The last two columns of table 4 use the change in average hourly earnings 4 quarters in the past or 4 quarters in the future as an instrumental variable for the change in average hourly earnings.[31] This instrument is more strongly correlated with the change in average hourly earnings, leading to a substantial reduction in the estimated standard errors associated with the wage elasticities, although little change in the point estimates. Again, the estimated coefficients of the selection terms are positive and statistically significant, but their inclusion leads to little change in the estimated wage elasticities.

Irrespective of the choice of instrumental variable, the estimates of the wage elasticity of hours per week, which in the context of the model can be interpreted as estimates of the intertemporal substitution elasticity, are in the same range as those obtained by MaCurdy (1981) and Altonji (1986) using annual hours data from the PSID. There is no evidence to suggest that analysis of quarterly observations on hours per week will lead to a new assessment of the size of the intertemporal substitution elasticity. Either the intertemporal labor-supply elasticity is relatively small (on the order of estimates already in the literature), or, contrary to the model in Section II, one cannot recover its magnitude from the curvature of weekly earnings schedules. Since there is considerable variation in weekly earnings and hours in the SIPP sample, it remains an interesting question how this variation is determined, if not by preferences over intertemporal labor supply.

IV. SUMMARY

This paper begins with a descriptive analysis of short-term hours variation among males in the Survey of Income and Program Participation. A unique feature of the SIPP is its 4-month interview schedule. This schedule permits a detailed investigation of the components of variance in individual labor supply over time. A simple decomposition suggests that

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[31] I tested whether changes in wages are correlated differently with changes 4 quarters in the past or 4 quarters in the future but found no significant difference.

quarter-to-quarter variation in hours worked arises from three sources: movements in and out of employment for an entire quarter; changes in the number of weeks worked, conditional on employment; and changes in the number of hours per week. For individuals with the same employer over the sample period, these three components contribute about equal shares to the overall variation in hours around individual-specific means. For others, movements in and out of employment for the entire quarter contribute the largest share of variance, whereas changes in hours per week contribute the smallest share.

Motivated by the observation that weekly labor supply exhibits some minimum hours threshold, the second part of the paper constructs a simple contracting model with mobility costs and a nonconvexity in the relation between hours per worker and output. This model generates a two-part hours schedule. In states of low demand, some individuals are temporarily laid off, while others are employed at a minimum weekly hours level. In states of high demand all individuals work hours in excess of the minimum threshold. The contract is supported by a simple schedule relating earnings per worker to hours per worker. Under the assumption of within-period additive separability between leisure and consumption, the implied elasticity between hours and average hourly earnings is the conventional intertemporal labor supply.

The model offers a simple explanation for the role of demand-side variables in an individual labor-supply function. According to the model, changes in weeks worked occur at fixed average hourly earnings, reflecting the realization of employment demand. Changes in annual measures of labor supply, which consist in part of changes in weeks worked, will therefore be correlated with measures of employment demand faced by an individual's employer, even after controlling for average hourly earnings.

The model implies that estimates of the intertemporal labor-supply elasticity can be obtained from estimates of the wage elasticity of hours per week for individuals observed in the same contract over time. Estimates are constructed from quarter-to-quarter changes in labor supply for men with the same employer during the SIPP sample period. The estimation procedures take account of measurement error in average hourly earnings and possible selection biases associated with movements in and out of employment for an entire quarter. Even with these adjustments, the estimates of the wage elasticity of hours per week are relatively small, ranging from .10 to .22.

If the basic point of the model is correct, however, it is important

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to realize that variation in labor supply may be much more sensitive to changes in productivity than these estimates suggest. This is because changes in productivity have a direct effect on the fraction of individuals employed at a given firm and on the number of weeks worked by a given individual. The evidence in this paper shows that changes in weeks are a major source of variation in individual labor supply and that these changes are highly correlated with industry-demand conditions. However, there is little indication that these changes occur along a conventional labor-supply schedule.

DATA APPENDIX

The data from consecutive interviews of the SIPP are released as cross-sectional data sets. A panel can be constructed by merging the information from the 9 interview "waves." Individuals are identified in each wave by a combination of three variables: the sample unit identifier, the entry address identifier, and the person number. I merged information for men whose age is between 22 and 62 in all waves of the 1984 panel by these three identifiers. As noted in the SIPP User's Guide (U.S. Department of Commerce (1987), chapter 6), it is possible that some individuals are incorrectly merged, since the SIPP interviewers occasionally assign the same series of identifiers to different people. This is most likely to happen in the event of a marital dissolution. On the other hand, individuals who move cannot be merged because the SIPP sampling frame is based on residential location. The SIPP also adds new members to the sample as they move into existing sample households. I found 8280 individuals who reported information in 8 consecutive interviews in the 1984 SIPP panel and another 8458 individuals for whom information was missing in at least one interview. The following table compares the characteristics of those with complete longitudinal data and those without:

Comparisons of Individuals with Complete and Incomplete Data

| Characteristic | Mean In Sample With: | |
(at first interview)	Complete Data	Incomplete Data
1. Age (years)	38.1	40.1
2. Education (years)	16.9	16.2
3. Percent worked 4 weeks in previous month	83.9	75.5
4. Monthly earnings ($) (including zeros)	1388.3	1208.9
5. Percent Black	8.4	10.8

A second significant reduction in the size of the sample occurs when individuals with imputed data on earnings, weeks worked, and hours per week are excluded. Of the 8280 individuals in the sample with complete longitudinal data, 1965 have an imputed value for one of these variables at some point in the sample. The following table compares the characteristics of those with imputed data and those without:

Comparisons of Individuals With and Without Imputed Data

| Characteristic | Mean In Sample With: | |
	Imputed Data	No Imputations
1. Age in 1986 (years)	40.4	40.3
2. Monthly earnings ($) excluding nonworkers:		
September 1983	1758.8	1791.4
December 1986	1955.3	1983.6
3. Usual hours per week excluding nonworkers:		
September 1983	42.8	42.4
December 1986	43.0	42.4
5. Percent Black	9.7	8.1

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