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4 Anticipated Unemployment, Temporary Layoffs, and Compensating Wage Differentials

John M. Abowd and Orley Ashenfelter

This paper models the competitive equilibrium wage rate when employment offers vary according to the amount of anticipated unemployment and unemployment risk. The competitive wage reflects a compensating differential which includes a certainty equivalent compensation proportional to the squared expected unemployment rate and a risk compensation proposal to the coefficient of unemployment variation. The factors of proportionality are half the inverse compensated labor supply elasticity and half the relative risk aversion, respectively. We use panel data to construct a model of anticipated unemployment and unemployment variance which depends on personal employment history and industry-wide and economy-wide factors. Compensating wage differentials ranging from less than one percent to more than fourteen percent are estimated for a two-digit industry classification over the years 1970–75.

4.1 Introduction

The appropriate theoretical framework for the interpretation of measurements of unemployment is once again a matter of controversy. The issues involved are important both because of the overwhelming role

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that measured unemployment plays in the discussion of public policies designed to mitigate it and because of the research strategies implied for understanding the causes and consequences of movements in it. One natural research strategy is to interpret unemployment as the equilibrium outcome of a worker's choices about job search or the intertemporal allocation of nonmarket time.[1] Of course, labor supply theoretic explanations of movements in unemployment are not complete without further specification, but this approach does imply a general strategy for the necessary research. An alternative approach is to treat unemployment as a constraint on individual behavior rather than a result of it.[2] Although this demand theoretic approach is also incomplete without further specification, it also implies a general strategy for further research.

Treating unemployment as a constraint on behavior may appear to have immediate normative implications for public policies regarding unemployment benefits or compensation. It may seem plausible that workers would, as a group, purchase insurance against an unpredictable exogenous constraint on the hours they are able to sell in the market. Whether this insurance is provided privately or governmentally would seem, then, to be a matter of form rather than substance. Predictable variations in the nature and extent of the risks of unemployment are surely extensive, so that a uniform governmental insurance benefit scheme would still leave considerable variation in the incidence of unemployment constraints to be compensated within the labor market. The purpose of this paper is to examine a theoretical model of market wage adjustments to compensate for the uninsured differences in the incidence of unemployment. We examine the empirical significance of these compensations using data from the Panel Study of Income Dynamics for the years 1967 to 1975 (Survey Research Center 1972ab 1973 1974 1975 1976).

Our analysis concerns two different but related issues. First, we extend the emerging empirical and theoretical analyses of the sources of compensating wage differentials to include a systematic treatment of the impact of constraints on hours at work.[3] Our analysis shows clearly how the determination of wage rates in the presence of fully anticipated constraints is systematically related to the determination of labor supply in the absence of such constraints, and demonstrates that these two issues are connected in a way that has not been fully appreciated in the past. The analysis provides a theoretical framework for measuring the impact of risky labor supply constraints on the determination of market wages. At the same time, our model also provides some evidence of how fruitful a research strategy that treats unemployment as a constraint on behavior is likely to be. We find that the labor market compensates anticipated layoffs and unemployment by 2–6 percent per year. The estimated compensations vary from industry to industry but are relatively stable from year to year within industries. Section 4.2 provides a theoretical analysis of equilibrium wage compensations. Section 4.3 disusses the empirical methodology and results.

4.2 A Theoretical Model of Compensating Differentials and Employment Constraints

In this section we consider the determination of wages when workers may choose employment in either of two sectors. In the unconstrained sector the worker may choose the optimum labor supply given the prevailing wage. In the constrained sector the worker accepts a fixed wage and employment conditions contract which sets the average number of work hours and employment variability. In a competitive equilibrium with identical workers, the utility of these two employment situations will be equal. This condition is used to characterize the relationship between the wage in the constrained sector and the conditions of employment. Specifically, we relate the equilibrium wage to the expected extent of unemployment and hours variability.

In the absence of a constraint on the hours a worker may sell in the market, the usual assumptions about preferences between nonmarket time and commodities lead to an optimal offer of labor supply $h^0(w, p, y)$ and optimal commodity demands of $x^0(w, p, y)$ that depend on the wage rate w, commodity prices p, and nonlabor income y. The indirect or maximal utility for this consumer-worker is than $V(w, p, y)$ which also depends on w, p, and y. Consider now a constraint $\bar{h} < h^0(w, p, y)$ which prevents the worker from selling more than $\bar{h}$ hours on the labor market. Faced with a binding constraint, the worker will supply $\bar{h}$ hours of work, demand commodities $x^*(\bar{h}, p, w\bar{h} + y)$, and achieve maximal utility of $V^*(\bar{h}, p, w\bar{h} + y)$. For the same w, p, and y the constraint on hours will, of course, lead to a maximal utility $V^*(\bar{h}, p, w\bar{h} + y)$ which is less than $V(w, p, y)$. The constrained worker would then reach a lower utility level than would be the case in the unconstrained job. The maximum of V^* with respect to $\bar{h}$ is simply $V(w, p, y)$ and is achieved at the point $\bar{h} = h^0$. V^* is graphed in figure 4.1 as a function of $\bar{h}$ for given w, p, and y. We assume V^* is concave in the relevant region although this is necessary only at the point $\bar{h} = h^0(w, p, y)$.

If the worker is faced with a choice of a job in the unconstrained sector versus a job in the constrained sector at the same wage rate, the worker would always prefer employment in the unconstrained sector. Labor market equilibrium will entail the condition

$$V^*(\bar{h}, p, w^*\bar{h} + y) = V(w, p, y) \tag{1}$$

where w is the prevailing wage in the unconstrained sector and w^* is the equilibrium wage rate when the employment contract requires labor

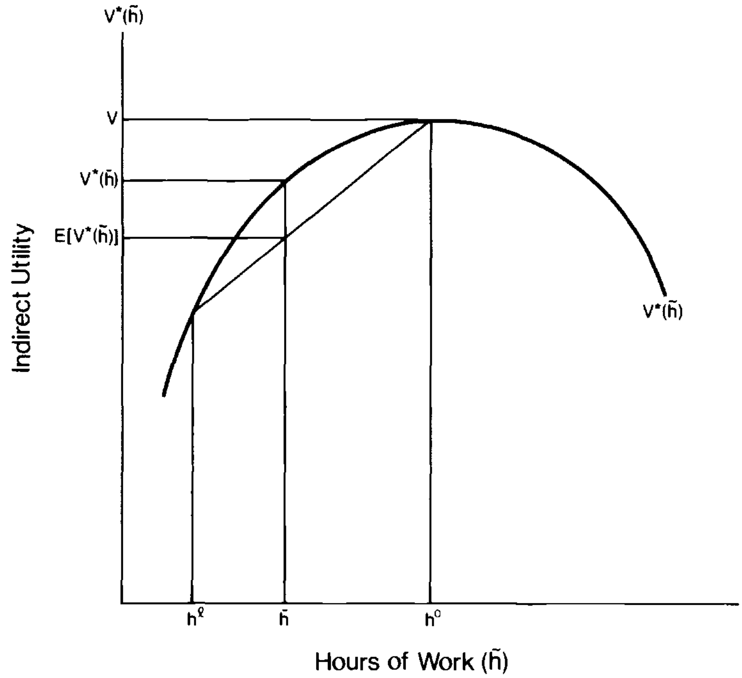

Figure 4.1 Sources of Compensating Differentials Illustrated Using Hours of Work and Constrained Maximal Utility

supply of $\bar{h}$ hours.[4] Equation 1 defines the proportionate compensating wage differential $(w^* - w)/w \approx \ell n(w^*/w)$ implicitly. Figure 4.1 demonstrates this idea for the choice between jobs that entail h^0 and $\bar{h}$ hours at work. The compensating wage differential is simply the increase in the wage rate sufficient to eliminate $V - V^*(\bar{h})$, the utility difference.[5] The consumer worker is indifferent between the wage-hours package w, h^0 and the wage-hours pair w^*, $\bar{h}$. Both allow the attaintment of the same utility level.[6] Formally, w^* is the solution of the implicit equation 1 and may be written as

$$(2) \qquad w^* = w^*(\bar{h}, w, p, y)$$

We will discuss some straightforward approximations to equation 2 below.

The preceding alalysis supposes that the hours of work constraint $\bar{h}$ is known with certainty. In some situations it may be more realistic to assume that $\tilde{h}$ is a random drawing from a distribution of possible hours constraints with expected value $\bar{h}$. The worker must choose between a job in the constrained sector which offers the known wage w^{**} and a known distribution of hours of work, versus a job in the unconstrained sector with the known combination w, h^0. Labor market equilibrium will then require that w^{**} compensate for the riskiness of fluctuations in hours worked and for the presence of the mean constraint on hours worked. Figure 4.1 also illustrates this point. We have assumed in the figure that $\tilde{h} = h^0$ with probability π, and $\tilde{h} = h^\ell$ with probability $1 - \pi$. The probability is chosen such that $E[\tilde{h}] \equiv \bar{h} = \pi h^0 + (1 - \pi)h^\ell$. Expected utility is then $E[V^*(\tilde{h})] = \pi V^*(h^0) + (1 - \pi)V^*(h^\ell)$. The wage difference $w^{**} - w^*$ compensates for the utility difference $E[V^*(\tilde{h})] - V^*(\bar{h})$, which is the incremental utility loss associated with the addition of risk to the model with only a mean constraint. As usual, the sign of the compensating differential depends on the worker's attitude toward risk. The case of risk aversion, represented by the concavity of the function $V^*(\bar{h})$ in figure 4.1 implies that additional risk adds to the compensating wage differential $w^{**} - w^*$. This may be seen by noticing that a mean preserving decrease in h^ℓ would lead to greater variance in hours, a greater (absolute) difference in $E[V^*(\tilde{h})] - V^*(\bar{h})$, and a greater compensating wage differential. The concavity of the function $V^*(\bar{h})$ is guaranteed in the neighborhood of $\bar{h} = h^0$ so that for small deviations of $\bar{h}$ from h^0, increased risk will generally imply an increased compensating wage differential.

In general, labor market equilibrium will now require that

$$V^{**}(w^{**}, p, y; \theta) = V^*(\bar{h}, p, w^*\bar{h} + y) \tag{3}$$

where $V^{**} = \int V^*(\tilde{h}, p, w^{**}\tilde{h} + y)f(\tilde{h}; \theta)\, d\tilde{h}$ is expected utility and θ represents the parameters of the density function for hours of work offered. Equation 3 defines w^{**} implicitly as

$$w^{**} = w^{**}(w^*, p, y; \theta) \tag{4}$$

which is the wage rate which must accompany an employment contract which offers expected constraints and uncertain employment.

The representation of a simplified unemployment insurance (UI) system does not add any major complications. We use the simple specification that the UI system replaces a fixed proportion γ of lost earnings. Labor income is, then, $w^{**}[\tilde{h} + \gamma(h^0 - \tilde{h})]$. The number of hours actually compensated is $\tilde{h} + \gamma(h^0 - \tilde{h})$, which we call effective hours.[7] Equilibrium condition 1 is replaced by

$$V^{\gamma *}(\bar{h}, p, w^*[\bar{h} + \gamma(h^0 - \bar{h})] + y) = V(w, p, y) \tag{1'}$$

Effective hours, and not actual hours, enter the relevant budget constraint. It should be noted that we used h^0 which depends on w and not w^* as the hours base for calculating lost earnings. The relevant comparison for determining the opportunity cost of the layoff or hours reduction is with the labor supply in the unconstrained sector. The wage w is the only wage at which the worker can optimize hours of work directly. The market operates to make the utility level provided by the insured constrained job equal to the utility level provided by $h^0(w, p, y)$. The utility level provided by $h^0(w^*, p, y)$ is irrelevant since no worker can achieve it. Condition 1′ introduces an important new twist into the analysis. Depending on the values of γ and $h^0 - \bar{h}$, it is possible for w^* to be less than w. This occurs whenever the utility payment for the UI benefits exceeds the utility loss from being required to consume too much leisure. The nature of this tradeoff is clear from the approximations below. When the hours offer $\tilde{h}$ is a random variable with mean $\bar{h}$ in the insured sector, the equilibrium condition 3 becomes

$$V^{\gamma**}(w^{**}, p, y; \theta) = V^{\gamma*}(\bar{h}, p, w^*[\bar{h} + \gamma(h^0 - \bar{h})] + y) \tag{3'}$$

The compensating wage differentials defined implicitly by (1′) and (3′) will, of course, depend on γ as well as the wage, price, nonlabor income triple w, p, y, and the θ parameters.[8]

We now address the problem of deriving useful approximations to the compensations expressed in equations 2 and 4. The results reveal that the parameters of the approximations have familiar interpretations which relate to conventional labor supply and risk analysis. We combine the approximations to form an estimating equation for the wage-generating function for a panel of continuously employed (or temporarily laid off) workers.

Let $U(x, 1-h)$ be a strictly quasiconcave, twice continuously differentiable utility function, and let $px = wh + y$ be the budget constraint. It is convenient[9] to work with the minimum expenditure function $R(w, p, v)$, defined as

$$R(w, p, v) = \min_{\{x,\, h\}} px - wh \tag{5}$$

subject to $v = U(x, 1-h)$ and $0 < h < 1$. In the absence of unemployment insurance the minimum expenditure function for an individual constrained by $h = \bar{h}$ with certainty is given by

$$R^*(\bar{h}, p, v) - w\bar{h} = \min_{\{x\}} px - w\bar{h} \tag{6}$$

subject to $v = U(x, 1-\bar{h})$. Equilibrium condition 1 is equivalent to

$$R^*(\bar{h}, p, v) - w^*\bar{h} = R(w, p, v) \tag{7}$$

Expanding the left-hand side of (7) around the point $\bar{h} = h^c(w, p, v)$, the compensated labor supply function, yields the second order approximation

$$\frac{w^* - w}{w} \approx \frac{1}{2}\frac{1}{e}\frac{[\bar{h} - h^0(w, p, y)]^2}{\bar{h}h^0(w, p, y)} \tag{8}$$

where e is defined as the compensated labor supply elasticity for the unconstrained individual. The compensating differential is approximately proportional to the squared expected unemployment rate $u \equiv (h^0 - \bar{h})/h^0$. The factor of proportionality is half the inverse compensated labor supply elasticity. The more inelastic the compensated labor supply schedule, the greater the compensating differential will be for anticipated underemployment or anticipated spells of unemployment. The formula is symmetric. It also determines the overtime premium for anticipated overemployment although equation 8 expresses the overtime premium as if it applied to the average rather than the marginal wage rate.

Next consider the introduction of a UI system which pays benefits as γ percent of lost earnings $w^*(h^0 - \bar{h})$. Replacing constrained hours $\bar{h}$ by effective hours in the budget constraint implies that the constrained minimum expenditure function is

$$(6') \qquad \begin{aligned} &R^{\gamma *}(\bar{h}, p, v) - w\{\bar{h} + \gamma[h^c(w, p, v) - \bar{h}]\} \\ &= \min_{\{x\}} px - w\{\bar{h} + \gamma[h^c(w, p, v) - \bar{h}]\} \end{aligned}$$

subject to $v + U(x, 1 - \bar{h})$.[10] Replacing R^* in equation 7 with $R^{\gamma *}$ and expanding the left-hand side around $\bar{h} = h^c$ yields, to second order,

$$(8') \qquad \frac{w^* - w}{w} = \frac{-\gamma(h^0 - \bar{h})}{h^0 + \gamma(h^0 - \bar{h})} + \frac{1}{2}\frac{1}{e}\frac{(h^0 - \bar{h})^2}{h^0[\bar{h} + \gamma(h^0 - \bar{h})]}$$

where $h^0 = h^0(w, p, y)$. The presence of the UI benefits causes the compensating differential to fall in proportion to the expected unemployment rate. The factor of proportionality is exactly the UI system replacement rate. Equation 8′ shows explicitly the tradeoff between UI benefits and the labor supply compensation. The function describing the compensating differential is approximately quadratic in the expected unemployment rate. For small expected unemployment rates the compensating differential is negative. As the expected unemployment rate increases, the compensating differential falls to a (negative) minimum and then increases. The unemployment rate corresponding to the minimum compensating differential and the rate corresponding to a zero compensating differential depend on the values of γ and e as well as on the optimal and constrained labor supplies. Because (8′) is approximately quadratic in the expected unemployment rate, however, the basic dependence is on the product γe. It is approximately true that the minimum differential occurs at the unemployment rate $u = \gamma e$ and the two zero compensating differentials occur at unemployment rates of zero and γe.[11] These three points determine the quadratic approximation.

If the labor supply constraint $\bar{h}$ is drawn from a distribution with $E[\bar{h}] = \bar{h}$ and $Var[\bar{h}] = \sigma^2$ then, in the absence of UI insurance, equilibrium condition 3 can be expanded in a Taylor series around $\bar{h}$ to yield the approximation

$$(9) \qquad \frac{w^{**}-w^{*}}{w^{*}} \approx \frac{1}{2} r \frac{\sigma^2}{\bar{h}^2}$$

where r is the relative risk aversion function when $\tilde{h}$ is the sole source of randomness in V^*.[12] Equation 9 implies that the incremental compensating differential implied by positive hours variance is approximately proportional to the squared coefficient of variation in hours of work. The factor of proportionality is half the relative risk aversion, r, which must be positive, at least when $\bar{h}$ is near h^0. When a UI system is in effect, equation 9 is modified to be

$$(9') \qquad \frac{w^{**}-w^{*}}{w^{*}} = \frac{1}{2} r \frac{\sigma^2}{\bar{h}[\bar{h}+\gamma(h^0-\bar{h})]}$$

where h^0 is evaluated at $h^0(w, p, y)$ as in (8′) above.

Although we apply the approximations 8 and 9 to individual data using individual unemployment rates, the theory has some implications for the use of aggregate unemployment rates in wage determination equations as well. Two points should be remembered. First, the error in hours of work, $\tilde{h}-\bar{h}$, is not a determinant of the compensating differential. Consequently, the individual's realized unemployment rate is not relevant; only the predictable component $h^0-\bar{h}$ affects the wage. Using the aggregate unemployment rate in the individual's wage determination equation is justified only if this variable is the appropriate expected unemployment rate for the individual. Second, the use of the squared aggregate unemployment rate confounds two effects. The coefficient on the squared expected unemployment rate measures the inverse labor supply elasticity, while the coefficient on the expected squared error (squared coefficient of variation) measures the risk compensation. Unless the aggregate unemployment rate is decomposed into predicted and error components, these effects will be confounded.

The approximations in equations 8′ and 9′ make clear the relationship between our model and labor supply decisions when layoffs are correctly anticipated. Workers may enter into explicit contracts such as collective bargaining agreements or implicit contracts with particular employers in order to secure employment in the constrained sector. A temporary layoff is one tool which the employer may use to make the expected constraint a realized constraint. If temporary layoffs were the only form of hours reduction that an employer used, then the key parameters θ in the distribution of $\tilde{h}$ would be the layoff probability π; the conditional expected completed duration given a layoff δ; and the conditional variance of the layoff duration ϕ^2. The implied compensating wage differential is

$$\frac{w^{**}-w}{w} = -\gamma\frac{\pi\delta}{h^0+(\gamma-1)\pi\delta} + \frac{1}{2}\frac{1}{e}\frac{\pi^2\delta^2}{h^0[h^0+(\gamma-1)\pi\delta]} \tag{10}$$
$$+\frac{1}{2}r\frac{\pi[(1-\pi)\delta^2+\phi^2]}{(h^0-\pi\delta)[h^0+(\gamma-1)\pi\delta]}$$

Equation 10 forms the basis of our empirical analysis. It shows clearly that the theoretical constructs—expected constrained hours and hours variance—may be replaced with the parameters of the layoff structure associated with the job. The probability of temporary layoff, the conditional mean layoff duration, and conditional duration variance are all observable characteristics of a given job, in principle. Consequently, equation 10 implies that $\ell(w^{**}/w)$ can be specified in terms of employment conditions that are identifiable in many survey data sets. In addition, this formulation does not require a priori determination of which workers are employed in constrained labor supply situations. As the probability of experiencing a layoff approaches zero, the compensating differential also approaches zero. Hence, the layoff probability for every worker can be used as a continuous indicator of the extent to which labor supply constraints are an important aspect of the worker's job.

In this section we have shown that the appropriate compensating differential for employment situations involving anticipated spells of unemployment can be expressed in terms of familiar labor supply parameters. The job-specific variables affecting the size of the differential can be expressed in terms of quantifiable properties of the layoff incidence and layoff duration characteristics of the job. In the next section we calculate empirical analogues of these variables and estimate the compensating wage differentials.

4.3 Empirical Analysis of Compensating Wage Differentials

The wage function in equation 10 holds at the market level under the assumptions of section 4.2. Specifically, given data on individual opportunity wages w, expected unemployment $\pi\delta$, and unemployment variance $\pi[(1-\pi)\,\delta^2+\phi^2]$, we could estimate the parameters γ, e, and r directly from (10). There are no genuine issues of simultaneous determination as long as unobserved individual heterogeneity is confined to opportunity wages, unemployment incidence, and layoff duration.[13] The common utility function parameters e and r will still determine the market-equilibrium-compensating differential. On the other hand, if individuals vary substantially in either the value of the compensated labor supply elasticity or the relative risk aversion parameter, the market will induce a sort in which firms are matched with employees who supply the appropriate constrained hours and hours variance at minimum cost to the firm. For example, if e varies over the population while r does not, then individuals with the highest e will be sorted into the firms which use temporary layoffs the most—large $\pi\delta$. This combination provides the firm with its demanded employment flexibility at the lowest cost. Individuals with small e, however, are sorted into jobs with small expected unemployment, since these individuals require the largest compensating differentials. If the distribution of e is known, then, in principle, the market sorting of individuals into jobs according to $\pi\delta$ can be inferred directly. In practice, these equilibria are difficult to generate and analyze except in the simplest of cases.[14] We will confine our analysis to the model in which all workers share a common γ, e, and r but may differ systematically and permanently according to w, π, δ, and ϕ^2. That is, we will allow personal heterogeneity in the opportunity wages and job characteristics but not in the utility function. This preserves equation 10 as an operational form of the equilibrium without requiring that all individuals be observationally identical, given observed characteristics.

Panel data permit us to construct measures of the probability of layoff and layoff duration parameters which control for the individual's personal characteristics, past layoff durations, industry-specific past layoff durations, and economy-wide layoff histories. Although this information is probably inferior to the individual's own information, there is no doubt that controlling for personal history differences gives a substantially more accurate measure of the individual's employment situaton than we could get from annual survey averages produced separately for different industries and occupations. As in other empirical analyses of expectational equilibria, our task is to identify the consistent patterns in the employer's behavior and estimate the relationship between these forecastable actions and the market wage. We do observe that there are permanent differences across industries in the extent and duration of layoffs; however, there is also substantial intraindustry variability which reflects firm-specific and individual-specific differences in employment conditions. Our method of constructing π, δ, and ϕ^2 will capture these employer and individual components as well as industry-specific and economy wide differences in the unemployment expectations of the workers.

Formally, let d_{it} be the duration of unemployment for individual i in year t. Let z_{it} be the individual's employment history and personal characteristics up to year t, including d_{it-1}, d_{it-2}, d_{it-3} and economy histories. Consider the two equation systems for the latent variables y_{1it} and y_{2it}:

$$(11) \qquad \begin{aligned} y_{1it} &= z'_{it}\,\beta_1 + \varepsilon_{1it} \\ y_{2it} &+ z'_{it}\,\beta_2 + \varepsilon_{2it} \end{aligned}$$

where $\varepsilon_{it}(2 \times 1)$ is distributed $N(\mathbf{0}, \Sigma)$ independently and identically for all i and t, and β_1 and β_2 are unknown parameter vectors. Define the event

$$(12) \qquad \ell_{it} = \begin{cases} 1 \text{ if } i \text{ is laid off in year } t \\ 0 \text{ otherwise} \end{cases}$$

We use equation system 11 along with the following sample selection rules to determine the individual's employment-specific layoff parameters

$$(13) \qquad d_{it} = \begin{cases} y_{1it} \text{ if } \ell_{it} = 1 \\ 0 \text{ if } \ell_{it} = 0 \end{cases}$$

and

$$(14) \qquad \ell_{it} = \begin{cases} 1 \text{ if } y_{2it} > 0 \\ 0 \text{ otherwise} \end{cases}$$

Following Heckman (1979), the selection rules 13 and 14 imply that

$$(15) \qquad \pi_{it} \equiv \Pr\{\ell_{it} = 1\} = \int_{\frac{-z'_{it}\beta_2}{\sqrt{\sigma_{22}}}}^{\infty} f(s)ds$$

$$(16) \qquad \pi_{it}\delta_{it} \equiv E[d_{it} \mid z_{it}] = \pi_{it}(z'_{it}\beta_1 + \frac{\sigma_{12}}{\sqrt{\sigma_{22}}}\gamma_{it})$$

and

$$(17) \qquad \phi_{it}^2 = \sigma_{11}[(1 - \frac{\sigma_{12}^2}{\sigma_{11}\sigma_{22}}) + \frac{\sigma_{12}^2}{\sigma_{11}\sigma_{22}}(1 - \frac{z'_{it}\beta_2}{\sqrt{\sigma_{22}}}\lambda_{it} - \lambda_{it}^2)]$$

where $f(s)$ is the standard normal density, $F(s)$ is the cumulative normal distribution function, and

$$\lambda_{it} \equiv f(\frac{z'_{it}\beta_2}{\sqrt{\sigma_{22}}}) / F(\frac{z'_{it}\beta_2}{\sqrt{\sigma_{22}}})$$

Equation 15 models π_{it} as a probit function given the history contained in z_{it}. Equation 16 defines the unconditional mean duration as the expectation of the conditional mean durations δ_{it} and zero. Equation 17 defines the conditional variance ϕ_{it}^2. We use these quantities to form the unconditional variance $\pi_{it}[(1 - \pi_{it})\, \delta_{it}^2 + \phi_{it}^2]$. The statistical model of equations 15–17 provides an empirical counterpart to each of the required theoretical employment condition measures. We turn next to the estimation of the parameters of the layoff duration model and the compensating differentials.

4.3.1 The Data

Estimates are based on waves 1–9 of *The Panel Study of Income Dynamics,* (PSID) corresponding to calendar years 1967–75.[15] Only the 3,318 households in the wave 9 release of the probability sample were eligible for inclusion in the estimation sample. Since we used three years of employment history to calculate the unemployment measures, estimated differentials cover the period from 1970 (wave 4) to 1975 (wave 9). In order to be included in a particular year's estimation sample, an observation was required to come from a household with a white male head who was interviewed in the employed battery in the current year and each of the three previous years. This last requirement means that an individual must have been employed or temporarily laid off at the time of the interview (usually March of the following year) for four consecutive years. The unemployed battery is given to individuals who have experienced a prolonged layoff or who are unemployed and do not anticipate returning to their previous jobs. This selection rule, although primarily dictated by data availability considerations, results in an analysis sample of very stably employed individuals. On average, the sampled individuals have twenty-four years of labor force experience and over nine years of employer tenure. We are not concerned here with extending our conclusions to a population of less stably employed individuals such as youths or females. Rather, we realize that our conclusions apply primarily to individuals who have made relatively long-term employment commitments in industries and occupations which display substantial differences in short-term unemployment and temporary layoff patterns.

About forty percent of the 3,318 cases qualify for inclusion in any one year's estimation sample. Changes in the head of household and failure of the requirement of four consecutive interviews in the employed battery were about equally responsible for deletions. Experienced users of the PSID data will recognize that samples of 1,200–1,300 are common when continuous histories of individuals (as opposed to households) are drawn from the probability sample. Table 4.A.1 presents definitions and annual summary statistics for all variables used in the analysis.

4.3.2 Extimation and Measurement of the Unemployment Variables

The basic unemployment variable used in our analysis is $DSCRP_{it}$, which is defined as the number of hours individual i spent unemployed in year t.[16] Hours unemployed are measured as the product of days spent unemployed, or on strike, times the average hours worked per week when employed regularly. The variable $DUNEM_{it}$ is defined as one when $DSCRP > 0$ and zero otherwise. Table 4.1 presents the results of estimating a probit equation for DUNEM (column 1) and a selection-bias-corrected conditional expectation equation for DSCRP (column 2), given that layoff duration is positive.

The probit equation in column 1 of table 4.1 measures the effect of the individual's employment history on π_{it}. The determinants of π_{it} are the individual's history of layoff durations for the past three years (DSCRPM1–DSCRPM3), the industry-specific average layoff durations for the past three years (MDSCRPM1–MDSCRPM3), and the economy-

Table 4.1 **Estimates of the Parameters of the Layoff Probability and Conditional Duration Equations (Standard errors in parentheses)**

Variable	DUNEM[a] (Probit) (1)	DSCRP[b] (Lambda) (2)
DSCRPM1	1.813 (.100)	.384
DSCRPM2	.726 (.098)	.277
DSCRPM3	.492 (.083)	.131
MDSCRPM1	−3.848 (.749)	−1.395
MDSCRPM2	−2.031 (.879)	−.376
MDSCRPM3	−.986 (.684)	−.914
YDSCRPM1	.053 (2.537)	3.761
YDSCRPM2	−10.809 (3.737)	−3.284
YDSCRPM3	7.021 (2.205)	3.404
SCHCLASS1	.987 (.066)	−.014
SCHCLASS2	.654 (.063)	−.022
SCHCLASS3	.451 (.065)	.028
EXPERIENCE	−.010 (.007)	−.001
$\text{EXPERIENCE}^2/100$	.006 (.014)	.007
TENURE	−.028 (.008)	−.012
$\text{TENURE}^2/100$	.085 (.027)	.020
LAMBDA	n.a.	.135
Constant	−.871 (.176)	.117

Table 4.1 (continued)

Variable	DUNEM[a] (Probit) (1)	DSCRP[b] (Lambda) (2)
R^2	n.a.	.111
S. E. equation	n.a.	.314
ℓn likelihood	−2185.870	n.a.
χ^2 (16)	9698.044	n.a.

[a]Estimated by maximum likelihood. Equation also includes industry dummy variables.
[b]Estimated by OLS. Equation also includes industry dummy variables. The reported S. E. equation and R^2 have been corrected.

wide average layoff durations for the same period (YDSCRPM1–YDSCRPM3). Permanent interindustry differences are captured by a set of fifteen industry dummy variables. Personal characteristics are measured by schooling dummy variables (relative to college graduates) (SCHCLAS1–SCHCLAS3), total labor force experience (EXPERIENCE, EXPERIENCE^2), and total time spent with the same employer (TENURE, TENURE^2). Although the determinants of π_{it} are not the major concern of this analysis, it is worth noting that personal unemployment history is an important predictor of current unemployment probability, even when industry average durations and economy-wide durations have been held constant. We interpret this effect as measuring employer and individual elements of the employment contract. No attempt is made to separate the employer-specific effect from the individual-specific effect since we remove a person effect from the wage equation used later in the analysis. The effects of common influences on all employers in a given industry are captured by the industry-specific variables. These allow both permanent and serially correlated transitory effects of unemployment history to influence predicted layoff probabilities. It is interesting to note that lagged industry layoff durations decrease the probability of a current layoff, given the industry dummy variables. This indicates that workers (and employers) expect unemployment to undershoot the long-term average after an exceptionally bad year and to overshoot the long-term average after an exceptionally good year. Economy-wide influences are mixed and rather imprecisely estimated. Personal characteristics have reasonable effects. High school dropouts (SCHCLAS1) are the most like to suffer layoffs, followed by high school graduates (SCHCLAS2), college dropouts (SCHCLAS3) and college graduates. Increases in employment tenure substantially reduce layoff probabilities (at a decreasing rate); lifetime labor force experience has a weak negative effect on layoffs (also at a decreasing rate).

Column 2 of table 4.1 shows the selection-bias-corrected estimates of the conditional duration equation. The pattern of effects is similar to the layoff probability equation, but no assessment of statistical significance can be made from the least-squares estimates. The implied interequation correlation coefficient is $\sigma_{12}/\sqrt{\sigma_{11}\,\sigma_{22}} = .43$. Under the null hypothesis of $\sigma_{12}=0$, the standard error of the coefficient on LAMBDA is .181. Although the hypothesis that $\sigma_{12}=0$ cannot be rejected at conventional significance levels, we do not want to ignore a sample correlation of .43; hence, we maintain the selection bias correction when calculating δ_{it} and ϕ^2_{it}.

Table 4.2 provides a comprehensive summary of observed layoff incidence and duration as well as the layoff incidence, expected duration, and duration variance implied by the equations in table 4.1. Table 4.2 reveals the substantial and relatively permanent interindustry differences in the use of temporary layoffs as a means of reducing work hours for long-term employees.[17] The row labeled DUNEM shows the actual annual average layoff frequency for each industry. The row labeled FDUNEM is the predicted layoff probability π_{it}, based on table 4.1. The DSCRP row shows the average layoff duration in thousands of hours.[18] The row EXPDURAT shows the expected layoff duration $\pi_{it}\delta_{it}$, based on table 4.1. The unconditional variance based on table 4.1 is shown in the row VARDURAT. Durable manufacturing industries (30–33) and the construction industry (51) have the highest layoff incidences and expected durations. Nondurable manufacturing (40, 45) also shows substantial use of layoff unemployment. These same industries exhibit higher unemployment risk (measured by VARDURAT) also.[19] Government (92) and professional service industries (86, 87) use layoff unemployment relatively little. Overall, 1975 was the year with the highest predicted unconditional unemployment duration (47 hours) and the highest observed unconditional unemployment duration (64 hours). The year 1971 had the largest overprediction (10 hours) while 1975 had the largest underprediction (−17 hours).

4.3.3 Estimation of the Wage Equation

The empirical counterparts of the variables in equation 10 were defined using the predicted probabilities, durations, and variances estimated in table 4.1 and summarized in table 4.2. We used reported hours (HACT_{it}) plus unemployed hours (DSCRP_{it}) as our measure of desired hours (HOPT_{it}). The empirical counterparts of the expressions in (10) are:

$$\text{HEFF}(\gamma) \equiv h^0 + (\gamma - 1)\pi\delta = \text{HOPT}_{it} + (\gamma - 1)\text{EXPDURAT}_{it} \tag{18}$$

$$\text{CERTEG}(\gamma)_{it} \equiv \frac{-\pi\delta}{h^0 + (\gamma - 1)\pi\delta} = \frac{-\text{EXPDURAT}_{it}}{\text{HEFF}(\gamma)_{it}} \tag{19}$$

Table 4.2 **Unemployment, Expected Unemployment, and Compensating Differential Measures for Annual Industry Aggregates**

	1970	1971	1972	1973	1974	1975
1. DSCRP	.0017	.0558	.0235	.0402	.0802	.0595
EXPDURAT	.0325	.0532	.0327	.0455	.0496	.0567
VARDURAT	.0184	.0309	.0152	.0217	.0240	.0318
DUNEM	.0203	.1194	.0458	.0685	.1473	.1429
FDUNEM	.0910	.1079	.0705	.0906	.1114	.1160
Comp. Diff.	2.564	5.205	6.779	7.469	4.888	4.441
(Std. Error)	(.430)	(.961)	(1.592)	(2.232)	(1.072)	(.815)
30. DSCRP	.0095	.0522	.0638	.0278	.0596	.0573
EXPDURAT	.0332	.0537	.0208	.0442	.0501	.0394
VARDURAT	.0183	.0288	.0121	.0209	.0214	.0198
DUNEM	.1207	.1429	.2143	.0943	.2076	.1818
FDUNEM	.1320	.1539	.0963	.1337	.1475	.1428
Comp. Diff.	3.968	4.792	2.087	4.878	4.881	3.859
(Std. Error)	(.681)	(.846)	(.349)	(1.130)	(1.130)	(.721)
31. DSCRP	.0097	.0375	.0149	.0267	.0117	.0593
EXPDURAT	.0204	.0355	.0166	.0307	.0349	.0329
VARDURAT	.0117	.0196	.0092	.0142	.0186	.0179
DUNEM	.0435	.0769	.0864	.0648	.0865	.1628
FDUNEM	.0859	.0999	.0624	.0891	.1045	.0965
Comp. Diff.	1.860	2.977	1.748	3.454	3.260	4.871
(Std. Error)	(.311)	(.507)	(.308)	(.814)	(.584)	(.915)
32. DSCRP	.2073	.0747	.0360	.0584	.1062	.1292
EXPDURAT	.0909	.0744	.0669	.0615	.1177	.1118
VARDURAT	.0439	.0317	.0310	.0315	.0538	.0468
DUNEM	.4456	.2410	.1714	.3000	.4407	.3529
FDUNEM	.2851	.2262	.1795	.2511	.3388	.2853
Comp. Diff.	8.450	8.101	5.992	7.047	14.011	10.281
(Std. Error)	(1.621)	(1.933)	(1.288)	(1.313)	(2.744)	(2.317)

	1970	1971	1792	1973	1974	1975
55. DSCRP	.0758	.0558	.0517	.0436	.0604	.0456
EXPDURAT	.0364	.0588	.0265	.0407	.0576	.0653
VARDURAT	.0211	.0298	.0148	.0205	.0278	.0286
DUNEM	.3239	.2727	.1312	.1129	.2121	.1667
FDUNEM	.1725	.2001	.1171	.1685	.2079	.1967
Comp. Diff.	2.865	4.233	3.492	3.746	5.372	5.820
(Std. Error)	(.507)	(.783)	(.678)	(.787)	(1.125)	(1.386)
57. DSCRP	.0050	.0395	.0088	.0025	.0075	.0383
EXPDURAT	.0126	.0219	.0167	.0226	.0190	.0224
VARDURAT	.0079	.0124	.0065	.0120	.0112	.0126
DUNEM	.0862	.1452	.0400	.0208	.0417	.0208
FDUNEM	.0661	.0761	.0520	.0700	.0811	.0800
Comp. Diff.	1.354	1.992	1.818	2.215	1.786	1.908
(Std. Error)	(.221)	(.329)	(.510)	(.458)	(.303)	(.318)
61. DSCRP	.0133	.0159	.0149	.0128	.0242	.0309
EXPDURAT	.0153	.0252	.0113	.0205	.0258	.0318
VARDURAT	.0090	.0140	.0066	.0112	.0141	.0158
DUNEM	.0674	.0315	.0516	.0446	.0828	.0898
FDUNEM	.0678	.0802	.0504	.0753	.0874	.0896
Comp. Diff.	1.672	2.042	.812	2.433	2.026	2.625
(Std. Error)	(.287)	(.343)	(.134)	(.451)	(.344)	(.511)
81. DSCRP	.0107	.0000	.0057	.0413	.0248	.0637
EXPDURAT	.0180	.0288	.0117	.0231	.0257	.0376
VARDURAT	.0103	.0161	.0068	.0129	.0142	.0163
DUNEM	.0333	.0000	.0286	.1316	.0811	.1471
FDUNEM	.0691	.0843	.0493	.0774	.0853	.0870
Comp. Diff.	1.327	2.075	.964	2.094	2.200	3.337
(Std. Error)	(.218)	(.348)	(.158)	(.348)	(.370)	(.786)

Table 4.2 (continued)

	1970	1971	1972	1973	1974	1975
33. DSCRP	.0161	.0329	.0097	.0197	.0374	.0782
EXPDURAT	.0191	.0351	.0166	.0286	.0357	.0354
VARDURAT	.0112	.0179	.0087	.0158	.0177	.0187
DUNEM	.0870	.1154	.0571	.0308	.1967	.1667
FDUNEM	.0852	.1056	.0702	.0990	.1165	.1086
Comp. Diff.	1.964	3.134	1.450	2.519	3.120	3.381
(Std. Error)	(.327)	(.591)	(.268)	(.435)	(.637)	(.616)
40. DSCRP	.0095	.0176	.0034	.0241	.0117	.0310
EXPDURAT	.0093	.0220	.0055	.0127	.0187	.0225
VARDURAT	.0104	.0143	.0070	.0111	.0138	.0146
DUNEM	.0588	.2051	.1290	.1667	.1250	.1026
FDUNEM	.1121	.1256	.0766	.1151	.1378	.1291
Comp. Diff.	1.509	2.438	1.151	1.879	2.269	4.330
(Std. Error)	(.245)	(.403)	(.187)	(.306)	(.398)	(.804)
45. DSCRP	.0399	.0188	.0049	.0351	.0465	.1261
EXPDURAT	.0372	.0367	.0258	.0429	.0580	.0562
VARDURAT	.0205	.0196	.0143	.0189	.0234	.0244
DUNEM	.2118	.0488	.0241	.1333	.1487	.2647
FDUNEM	.1326	.1186	.0893	.1296	.1459	.1423
Comp. Diff.	3.330	3.607	2.150	4.052	6.019	6.305
(Std. Error)	(.564)	(.626)	(.364)	(1.057)	(1.683)	(1.569)
51. DSCRP	.1226	.0427	.0411	.0574	.1072	.1982
EXPDURAT	.0604	.0835	.0537	.0725	.0938	.1006
VARDURAT	.0257	.0348	.0255	.0317	.0402	.0450
DUNEM	.2761	.1667	.1683	.1682	.2789	.4078
FDUNEM	.1781	.2095	.1594	.2061	.2360	.2356
Comp. Diff.	5.824	9.277	5.721	7.065	10.645	10.673
(Std. Error)	(1.379)	(2.310)	(1.226)	(1.613)	(2.619)	(2.437)

	1970	1971	1972	1973	1974	1975
85. DSCRP	.0162	.0047	.0177	.0204	.0353	.0319
EXPDURAT	.0167	.0340	.0153	.0230	.0286	.0353
VARDURAT	.0095	.0150	.0086	.0128	.0148	.0184
DUNEM	.0676	.0333	.0769	.0526	.0794	.0625
FDUNEM	.0653	.0794	.0519	.0759	.0892	.0895
Comp. Diff.	1.801	3.854	1.635	1.925	3.887	3.739
(Std. Error)	(.299)	(.872)	(.285)	(.323)	(.689)	(.707)
86. DSCRP	.0119	.0000	.0025	.0000	.0131	.0190
EXPDURAT	.0120	.0204	.0119	.0145	.0211	.0291
VARDURAT	.0072	.0114	.0063	.0084	.0118	.0158
DUNEM	.0351	.0000	.0208	.0000	.0600	.0328
FDUNEM	.0559	.0646	.0408	.0613	.0740	.0752
Comp. Diff.	1.303	1.817	1.382	1.170	1.907	3.392
(Std. Error)	(.213)	(.302)	(.279)	(.194)	(.318)	(.657)
87. DSCRP	.0011	.0189	.0006	.0002	.0040	.0216
EXPDURAT	.0104	.0219	.0070	.0155	.0216	.0220
VARDURAT	.0074	.0122	.0053	.0095	.0124	.0124
DUNEM	.0230	.0460	.0116	.0130	.0274	.0811
FDUNEM	.0697	.0812	.0503	.0792	.0915	.0870
Comp. Diff.	1.351	2.049	.944	1.473	2.368	2.248
(Std. Error)	(.227)	(.347)	(.155)	(.241)	(.394)	(.372)
92. DSCRP	.0176	.0022	.0252	.0176	.0035	.0050
EXPDURAT	.0160	.0197	.0116	.0160	.0203	.0196
VARDURAT	.0081	.0109	.0057	.0084	.0113	.0109
DUNEM	.0652	.0274	.0225	.0538	.0194	.0206
FDUNEM	.0502	.0564	.0346	.0529	.0624	.0601
Comp. Diff.	1.862	1.627	3.797	1.509	5.182	2.057
(Std. Error)	(.377)	(.274)	(.921)	(.266)	(.891)	(.344)

$$\text{(20)} \qquad \text{CERTSQG}(\gamma)_{it} \equiv \frac{.5\pi^2\delta^2}{h^0[h^0 + (\gamma - 1)\pi\delta]}$$

$$\equiv \frac{.5\ (\text{EXPDURAT}_{it})^2}{\text{HOPT}_{it}\ \text{HEFF}(\gamma)_{it}}$$

$$\text{(21)} \qquad \text{COEFVSQG}(\gamma)_{it} \equiv \frac{.5\pi[(1-\pi)\delta^2 + \phi^2}{(h^0 - \pi\delta)[h^0 + (\gamma - 1)\pi\delta]}$$

$$\equiv \frac{.5\ \text{VARDURAT}_{it}}{(\text{HOPT}_{it} - \text{EXPDURAT}_{it})\text{HEFF}(\gamma)_{it}}$$

From the definitions in equations 18–21 it is clear that there is a nonlinear dependence on γ (the UI replacement rate) in all three of the expressions which enter the equation for $\ell n(w^{**})$. We deal with this problem in two different ways. First, we estimate γ along with e, r, and the determinants of $\ell n(w)$ by nonlinear methods. Second, we compute the sample average UI replacement rates by industry for each year (MCOMPRAT_{it}) and use these replacement rates instead of γ in the formation of the effective hours variable HEFF(γ).[20]

Table 4.3 reports the results of direct estimation of γ, e, and r by nonlinear least squares. Columns 1 and 2 report the use of the reported hourly wage ℓn(WAGE) as the dependent variable and columns 3 and 4 use the calculated hourly wage ℓn(EARN/HACT) as the dependent variable. We use as variables determining the opportunity wage—last year's wage (LAGW)—collective bargaining status (UNION), schooling, labor force experience, tenure, region, and year.[21] Columns 1 and 3 show the results from the specification implied by equation 10. Columns 2 and 4 show the results from the specifications implied by equation 10 but allowing unanticipated unemployment (UNANTIC) expressed as a percentage of desired hours to enter the equation.[22] The first three rows of table 4.3 show the implied estimates of γ, $1/e$, and r. The row labeled $\hat{e}_{\text{MELO}}$ shows the implied estimate of e using Zellner's (1978) method.[23] The row labeled average compensating differential shows the percentage compensating differential gross of the implicit UI purchase payment.

In general, the results of columns 1 and 2 are quite favorable to our model. The implied ninety percent confidence interval estimate of $.093 \pm .085$ for labor supply elasticity is rather imprecise but certainly consistent with previously reported values for white males in both magnitude and precision of estimation.[24] The estimated γ is too large to be a UI replacement rate; however, we will discuss this problem in more detail below. The estimated relative risk aversion of 14.048 (± 3.714) reveals substantial distaste for unemployment risk. This result supports the hypothesis that the difference in risk aversion between employees and employers results in a demand for implicit labor contracts of the type developed here.[25] Overall, the compensating differential is 3.82 percent (± 1.26%) which is not an unreasonably large price for employers to pay for the right to set π, δ, and ϕ^2. Of this differential, .60 percent (± .55%)

Table 4.3 **Nonlinear Least-Squares Estimates of Wage Parameters under Alternative Specifications (Standard errors in parentheses)**

Variable Parameter	ℓn (WAGE)[a] (1)	ℓn (WAGE)[a] (2)	ℓn (EARN/HACT)[a] (3)	ℓn (EARN/HACT)[a] (4)
CERTEQG[b] γ	2.486	2.485	4.035	4.312
	(.530)	(.531)	(.600)	(.618)
CERTSQG l/e	8.107	8.102	7.283	9.234
	(4.572)	(4.592)	(5.292)	(5.624)
COEFVSQG r	14.048	14.044	28.177	29.555
	(2.258)	(2.263)	(3.073)	(3.171)
UNANTIC		−.001		.241
		(.058)		(.059)
LAGW	.767	.767	.772	.773
	(.009)	(.009)	(.009)	(.009)
UNION	.029	.030	.042	.041
	(.009)	(.009)	(.009)	(.009)
SCHOOL	.015	.015	.015	.015
	(.002)	(.002)	(.002)	(.002)
EXPERIENCE	.006	.006	.007	.007
	(.001)	(.001)	(.001)	(.001)
EXPERIENCE^2	−.013	−.013	−.014	−.015
/100	(.003)	(.003)	(.003)	(.003)
TENURE	.003	.003	.003	.004
	(.001)	(.001)	(.001)	(.002)
TENURE^2	−.002	−.002	−.004	−.004
/100	(.005)	(.005)	(.005)	(.005)
Average compensating	3.820	3.820	7.000	7.460
differential (%)[c]	(.767)	(.768)	(.995)	(1.041)
$\hat{e}_{\text{MELO}}$	.093	.093	.090	.079
	(.052)	(.053)	(.065)	(.048)
R^2	.693	.693	.699	.700
S. E. equation	.075	.075	.076	.076
Residual d.f.	5551	5550	5551	5550

[a]Regressions include annual regional dummy variables.
[b]This variable is defined as the negative of the theoretical variable so that its coefficient is γ and not $-\gamma$.
[c]Computed at the sample averages of CERTSQG and COEFVSQG, .00074 and .002293, respectively.

is attributable to the certainty equivalent compensation arising from the mean constraint, while 3.22 percent (±.85%) is attributable to risk compensation arising from the variability of unemployment, given the mean constraint. The addition of the unanticipated part of unemployment to the equation (UNANTIC) has no effect on the results.

The rows labeled Comp. Diff. and (Std. Error) in table 4.2 show the industry by year-estimated compensating differentials and associated standard error implied by column 1 of table 4.3. The reported compensating differentials are gross of the implicit payment for UI. The interindustry patterns reveal that automobile workers (32) received compensating differentials varying from 6.00 percent (1972) to 14.01 percent (1974) of hourly wages in exchange for bearing the mean constraints and employment risks discussed above. Construction workers (51) received compensating differentials ranging from 5.72 percent (1972) to 10.67 percent (1975). Workers in the public sector (92), however, received substantially smaller compensations ranging from 1.51 percent (1973) to 5.18 percent (1974). Other interindustry and time series comparisons are obvious from the results reported in table 4.2.

The results in table 4.3, columns 3 and 4, show that the parameter estimates are sensitive to the choice of dependent variable. The general pattern of results remains the same except for two points. First, the estimated effects are larger when calculated wage rates are used. The implied average compensating differential becomes 7.00 percent (±1.64%). Second, the addition of unanticipated unemployment to the equation using calculated wages as the dependent variable (column 4) does produce a substantial effect of .241 (±.097). This suggests that directly measured wage rates (WAGE) are probably more reliable measures of the hourly compensation relevant to implicit or explicit labor contracts than are the ex post calculated wage rates.

Table 4.4 reports estimates based on our second solution to the nonlinear dependence on γ. The compensating differential variables have been calculated using the observed industry average values of UI replacement rates (MCOMPRAT) reported in table 4.A.1. Using these variables we can remove a person-specific effect from each individual's opportunity wage by standard fixed effect methods. The results are presented in columns 1 and 2 for reported wages and columns 3 and 4 for calculated wages. Once again, the results for reported wages are highly supportive of the model. The estimated labor supply elasticity is .224 (±.729), which is more imprecise than the result in table 4.3 but still consistent with the previous estimates. The estimated risk aversion parameter, 8.982 (±1.828), is smaller in the fixed effect model. The estimated UI replacement rate $\hat{\gamma} = 1.206$ (±.681) now includes reasonable values in its ninety percent confidence interval. The average compensating differential is 2.14 percent (±.55%), which is smaller than the values

Table 4.4 Fixed Effect Estimates of the Compensating Differential Parameters (Standard errors in parentheses)

Variable Parameter	*ℓn* (WAGE)[a] (1)	*ℓn* (WAGE)[a] (2)	*ℓn* (EARN/HACT)[a] (3)	*ℓn* (EARN/HACT)[a] (4)
CERTEQG[b] γ	1.206	1.082	.718	.184
	(.414)	(.426)	(.413)	(.423)
CERTSQG l/e	1.140	.806	−2.327	−3.758
	(2.258)	(2.272)	(2.250)	(2.257)
COEFVSQG r	8.982	9.041	10.523	10.782
	(1.111)	(1.112)	(1.109)	(1.106)
UNANTIC		.083		.357
		(.066)		(.065)
Average compensating differential (%)[c]	2.144	2.133	2.241	2.194
	(.334)	(.334)	(.333)	(.332)
$\hat{e}_{MELO}$	.224	.139	[d]	[d]
	(.443)	(.391)		
R^2	.856	.859	.865	.866
S. E. equation	.223	.223	.222	.221
Residual d.f.	3863	3862	3863	3862

[a]Regressions include all variables in table 4.3.
[b]This variable is defined as the negative of the theoretical variable so that its coefficient is γ and not $-\gamma$.
[c]Computed as in footnote c, table 4.3.
[d]Coefficient of CERTSQG has the wrong sign.

in table 4.3. Finally, the addition of UNANTIC has a larger effect in column 2 than the comparable effect in table 4.3, column 2, but the effect is imprecisely estimated—.083 (± .109)—and its presence does not change any of the other effects substantially. As in the nonlinear model, the results are sensitive to the specification of the dependent variable. Columns 3 and 4 show that calculated wages give wrong-signed estimates of e and depend importantly on unanticipated unemployment. This is consistent with the conclusions for calculated wages shown in table 4.3.

Finally, we consider the implied wage payments for UI benefits. According to equation 10 and definition 19, the appropriate price is γCERTEG(γ). This price implies that wages fall by exactly the expected UI benefits in percentage terms. Table 4.5 shows the annual lower and upper 90 percent confidence bounds for the implicit UI price based on table 4.3, column 1, and data for CERTEQG found in table 4.A.1. It also shows the unconditional sample average UI benefits as a percentage of labor earnings. As we noted above, the estimated γ is too large to reflect the UI replacement rate parameter. Nevertheless, the implied wage payments are quite close to the observed unconditional averages con-

Table 4.5 **Annual Average UI Benefits Compared with Benefit Expectations Implied by Estimates of γ**

	1970	1971	1972	1973	1974	1975	Overall
Lower bound (%)[a]	2.30	3.18	1.97	2.75	3.45	3.62	2.87
Upper bound (%)[a]	4.78	6.62	4.11	5.72	7.17	7.53	5.97
Observed average (%)[b]	1.06	2.47	.72	1.10	1.58	2.38	1.54

[a]Lower and upper bounds of 90% confidence interval computed using $\hat{\gamma} = 2.486$ (table 4.3, column 1) and standard error of $\gamma = .530$. Values of CERTEQG are in table 4.A.1.
[b]Annual average UI benefits expressed as a percentage of total labor earnings (EARN + benefits) for the whole sample (MCOMPRAT · DUNEM).

sidering the fact that the actual UI benefits were not ever used in the analysis leading to the estimates in table 4.3. Furthermore, an estimated γ of about 1 would give values for the implicit price of UI which are fully consistent with the observed percentages. This is, essentially, the value of γ produced by the estimation technique used in table 4.4.[26] It is reasonable to conclude that our high estimates of γ give an approximate consistency with the competitive requirement that the wage in the constrained job should fall by the expected UI benefit level.

4.3.4 Conclusions

The empirical analysis supports our major theoretical points. Estimated compensating differentials range from less than one percent in industries where the workers experience little anticipated unemployment to over fourteen percent in industries which experience substantial anticipated unemployment and unemployment risk. We find the implied estimate of the compensated labor supply elasticity is around .09, which is consistent with the prevailing evidence on this parameter. Our estimate of a relative risk aversion parameter of 14 is high, but, when coupled with the low values of the coefficient of variation for unemployment, the implied risk premium is modest. Finally, we present some evidence that the implicit price of UI implied by our model is somewhat large under one estimation method and approximately equal to the expected UI benefits under the alternative estimation method.

Appendix

Table 4.A.1 contains variable definitions and summary statistics for all variables used in the analysis. All data manipulations and estimation were performed using the *Statistical Analysis System* (1979 version).

Table 4.A.1 **Definitions and Means for All Variables Used in the Empirical Analysis (Standard deviations in parentheses)**

Variable	Definition
WAGE	Reported wage in current \$/hr for main job
ℓn(WAGE)	Natural logarithm of reported wage
EARN	Reported total labor income in current \$
ℓn(EARN/HACT)	Natural logarithm of EARN/HACT
HACT	Reported average hours per week × reported weeks worked (thousands of hours)
HOPT	HACT plus hours unemployed and hours on strike (thousands of hours)
DSCRP	HOPT-HACT
DUNEM	Equals 1 if DSCRP > 0, equals 0 otherwise
FDUNEM	Predicted probability of layoff spell
CONDDURA	Equals DSCRP if DSCRP > 0, missing otherwise (Summary statistics are conditional sample means and standard deviations)
EXPDURAT	Predicted unemployment using the formula in the text
VARDURAT	Predicted variance of unemployment using the formula in the text
HBAR	HOPT-EXPDURAT
MCOMPRAT	Industry average unemployment benefits received as a percentage of lost labor earnings
HEFF	HBAR + MCOMPRAT × EXPDURAT (Note: in table 4.3 HEFF is evaluated at the estimate of γ, not at MCOMPRAT)
CERTEQG	− EXPDURAT/HEFF (expected percentage unemployment, defined as the negative of the theoretical quantity)
CERTSQG	.5(EXPDURAT^2)/(HOPT × HEFF) (Half of squared expected percentage unemployment. Table entries × 10^{-3})
COEFVSQG	.5 VARDURAT/(HBAR × HEFF) (Half of the squared coefficient of variation of unemployment duration. Table entries × 10^{-3})

1970	1971	1972	1973	1974	1975	Overall
5.146 (2.952)	5.456 (3.313)	5.764 (3.425)	6.036 (3.187)	6.478 (3.531)	7.216 (3.862)	5.992 (3.447)
1.517 (.481)	1.571 (.488)	1.624 (.494)	1.681 (.479)	1.750 (.484)	1.858 (.492)	1.663 (.499)
11501 (5932)	12121 (6493)	12917 (6980)	13662 (7305)	14403 (7828)	15155 (8293)	13243 (7261)
8.423 (.488)	8.477 (.488)	8.528 (.501)	8.581 (.501)	8.657 (.501)	8.716 (.508)	8.560 (.509)
2.329 (.610)	2.311 (.558)	2.319 (.551)	2.333 (.558)	2.279 (.548)	2.259 (.563)	2.306 (.567)
2.369 (.583)	2.343 (.538)	2.339 (.537)	2.360 (.537)	2.321 (.524)	2.316 (.525)	2.342 (.542)
0.040 (.46)	0.032 (.136)	0.021 (.110)	0.027 (.127)	0.042 (.157)	0.064 (.207)	0.038 (.150)
0.129	0.105	0.074	0.085	0.136	0.153	0.114
0.111 (.093)	0.120 (.110)	0.079 (.096)	0.108 (.108)	0.129 (.122)	0.126 (.124)	0.112 (.111)
0.312 (.283)	0.309 (.303)	0.279 (.303)	0.325 (.308)	0.307 (.317)	0.416 (.365)	0.331 (.319)
0.030 (.048)	0.042 (.060)	0.024 (.053)	0.034 (.069)	0.043 (.071)	0.047 (.073)	0.036 (.063)
0.016 (.016)	0.021 (.020)	0.012 (.020)	0.017 (.020)	0.021 (.023)	0.023 (.024)	0.018 (.021)
2.339 (.587)	2.302 (.541)	2.316 (.543)	2.327 (.544)	2.277 (.530)	2.269 (.528)	2.306 (.547)
0.082 (.087)	0.236 (.414)	0.097 (.096)	0.130 (.211)	0.116 (.010)	0.156 (.091)	0.136 (.211)
2.342 (.586)	2.311 (.539)	2.319 (.542)	2.330 (.543)	2.284 (.528)	2.278 (.528)	2.312 (.546)
.0142 (.0246)	.0197 (.0365)	.0122 (.0391)	.0170 (.0538)	.0214 (.0431)	.0224 (.0404)	.0178 (.0404)
.352 (2.526)	.703 (5.310)	.644 (5.733)	.970 (12.879)	.927 (6.516)	.902 (4.671)	.740 (6.966)
1.866 (2.872)	2.451 (4.102)	1.759 (7.536)	2.120 (5.239)	2.835 (7.288)	2.779 (4.892)	2.293 (5.528)

Table 4.A.1 (continued)

Variable	Definition
SCHOOL	Years of school (SCHCLAS1 < 12, SCHCLAS2 = 12, SCHCLAS3 > 12 and/or S < 16, SCHCLAS4 > 16)
EXPERIENCE	Years of actual labor force experience
TENURE	Years employed with the same employer
UNION	Equals 1 if covered by a collective bargaining agreement (1975) or member of a labor union (1970–74)
UNANTIC	(DSCRP–EXPDURAT)/HOPT (Unanticipated unemployment as a proportion of desired hours)
DSCRPM1 DSCRPM2 DSCRPM3	Lagged values of individual DSCRP; 1, 2, 3 years, respectively (Means only shown. Overall standard deviation = .130)
MDSCRPM1 MDSCRPM2 MDSCRPM3	Lagged values of industry average DSCRP; 1, 2, 3 years, respectively
YDSCRPM1 YDSCRPM2 YDSCRPM3	Lagged values of overall average DSCRP; 1, 2, 3 years, respectively
INDUSTRY	Aggregates of the PSID 2-digit industry classification (Sample percentages shown)
1	All industures not classified
30	Metal manufacturing
31	Machinery, including electrical, manufacturing
32	Motor vehicles, other transportation equipment manufacturing
33	Other durable manufacturing
40	Food and tobacco manufacturing
45	Other nondurable manufacturing
51	Construction
55	Transportation
57	Other public utilities
61	Retail, wholesale, other trade
81	Repair services
85	Business, personal, amusement services
86	Professional, noneducational services
87	Educational services
92	Government (nonmedical, noneducational)
n	

1970	1971	1972	1973	1974	1975	Overall
12.090	12.233	12.228	12.312	12.477	12.711	12.335
(3.157)	(3.112)	(3.131)	(3.083)	(3.004)	(2.870)	(3.069)
26.017	25.598	24.696	23.525	22.496	21.534	24.033
(11.204)	(11.207)	(11.502)	(11.024)	(11.396)	(11.580)	(11.429)
9.416	10.250	9.569	9.389	8.862	8.473	9.335
(8.540)	(8.990)	(8.564)	(8.728)	(8.702)	(9.140)	(8.791)
.356	.330	.325	.299	.297	.286	.317
.0049	−.0041	−.0017	−.0025	−.0007	.0083	.0007
(.0591)	(.0616)	(.0513)	(.0535)	(.0652)	(.0879)	(.0642)
.019	.042	.031	.025	.029	.042	.031
.026	.019	.041	.036	.030	.028	.029
.024	.027	.023	.045	.049	.043	.035
Same as above						
Same as above						
10.64	10.53	12.68	11.98	10.57	10.97	11.21
4.17	3.85	3.48	4.35	4.34	3.63	3.97
6.61	8.17	6.71	8.86	8.52	7.10	7.64
7.26	6.52	5.80	4.10	4.83	5.61	5.73
6.61	6.13	5.80	5.33	5.00	4.95	5.66
2.44	3.06	2.57	3.45	3.28	3.22	2.99
6.11	6.44	6.88	6.15	6.06	5.61	6.21
9.63	9.90	8.37	8.78	8.52	8.50	8.97
5.20	5.18	5.05	5.09	5.41	5.45	5.21
4.17	4.87	4.14	3.94	3.93	3.96	4.17
12.80	12.49	12.84	12.88	12.86	13.78	12.93
2.16	1.81	2.90	3.12	3.03	2.81	2.62
5.32	4.71	4.31	4.68	5.16	5.28	4.92
4.10	3.77	3.98	3.36	4.10	5.03	4.05
6.25	6.83	7.12	6.32	5.98	6.11	6.43
6.61	5.73	7.37	7.63	8.44	8.00	7.27
1,391	1,273	1,207	1,219	1,221	1,212	7,523

Notes

1. See Mortensen (1970) and Lucas and Rapping (1969) for examples of unemployment as an optimal labor supply strategy.

2. See Malinvaud (1977) and Ashenfelter (1980) for examples of unemployment viewed as a constraint on labor supply behavior.

3. See especially Rosen (1978 1974). Lewis (1969) provides an early theoretical analysis of the problem of joint worker-employer determination of work hours.

4. Other terms of employment may vary in order to force equality between V^* and V. For example, a lump sum payment could be made to the worker or the fringe benefit package could be altered. We assume that the full adjustment occurs in wage rates.

5. In figure 4.1, V^* depends only on $\bar{h}$ since w, p, and y are being held constant.

6. This condition of equilibrium follows from the same type of argument as Feldstein (1978 1976) uses in analyzing the effects of the unemployment insurance system on temporary layoffs.

7. Later we will use Taylor series approximations around h^0 to provide estimating formulas. Only the linear term $\gamma(h^0 - \bar{h})$ enters these formulas even if the UI system is modeled in a completely general way.

8. The wage w^{**} is an example of an implicit contract equilibrium wage in the spirit of Azariadis (1975) and Baily (1977). Condition 3′ ensures that the contract $w^{**}, f(\bar{h}; \theta)$, which leads to unemployment or underemployment is just as desirable as the full employment contract, w, h^0. The firm's production and demand conditions determine which type of contract it will offer.

9. It is easier to derive the relevant approximations using the minimum expenditure function rather than the indirect utility function when there are no risk problems. Of course, we will use the utility function when we consider risk.

10. For the appropriate value of $v = V(w, p, y)$, $h^c(w, p, v) = h^0(w, p, y)$. Hence, the UI system in equation 6′ is identical to the system underlying 1′.

11. The reader is reminded that the unemployment rate used here is the expected personal unemployment rate $u = (h^0 - \bar{h})/h^0$ and not the economy-wide unemployment rate.

12. The reader is referred to Abowd and Ashenfelter (1979) for the derivation of this result. The relative risk aversion function used here is a straightforward extension of the concept in Pratt (1964) and Arrow (1970).

13. Duncan and Stafford (1980) consider simultaneous determination of wages and compensating differentials in the context of union wages and working conditions. In this model, personal heterogeneity is assumed to influence the size of the required compensation.

14. Rosen (1974) works an example.

15. See Survey Research Center (1972a) for a description of the methods and sampling frame. See Survey Research Center (1972b 1973 1974 1975 1976) for the relevant questionnaires and variable definitions.

16. All annual hours variables, including DSCRP, are defined in thousands of hours.

17. Sherwin Rosen provided some valuable comments concerning the usefulness of identifying industry patterns when forecasting the hours of work constraint.

18. The average shown is unconditional so that it can be compared with EXPDURAT and VARDURAT. The data for DUNEM and the industry proportions in table 4.A.1 can be used to calculate the conditional average duration.

19. EXPDURAT and VARDURAT are not independent since $\pi\delta$ and $\pi[(1-\pi)\delta^2 + \phi^2]$ will, in general, be correlated when π, δ, and ϕ all depend on z_{it}. The sample correlation coefficient is .91.

20. PSID reports a variable: income from unemployment and workmen's compensation. The sample replacement rates were calculated as industry by year averages of the ratio of

this UI income variable to the implied lost earnings (WAGE*DSCRP) of the head. The average is taken only over nonzero values of DSCRP.

21. The use of last year's wage is intended to remove a person effect from the nonlinear regression. We do not mean to imply any dynamics. As is well known, a fixed person effect cannot be consistently estimated in a nonlinear regression model by the usual deviation from time average method.

22. Ken Wolpin suggested this variation.

23. The method minimizes the posterior expected loss from a Bayesian viewpoint and improves the mean squared estimate error from the sampling theory viewpoint. The method is designed for problems where the parameter of interest is the inverse of the estimated parameter.

24. Ashenfelter and Heckman (1974) estimated the ninety percent confidence interval for the compensated labor supply elasticity at .06 (± .05) at their sample means in their table 1. Keeley, et al. (1978a b) estimated the ninety percent confidence interval at .23 (± .17) at our sample means using their table 2.

25. See Azariadis (1975) and, especially, Baily (1974) for a discussion of the differential risk aversion hypothesis.

26. We have produced summary measures based exclusively on table 4.3, column 1, since it is, overall, the most reasonable equation. The ninety percent confidence interval for the UI price implied by the estimate in table 4.4, column 1, is 2.14 percent (± 1.21%) which includes all the observed values reported in table 4.5. We should also note, however, that Feldstein (1978) reports that reported UI benefits received by CPS respondents understate the aggregate payments by fifty percent. This supports our estimate of γ in table 4.3, column 1, summarized in table 4.5.

References

Abowd, J., and Ashenfelter, O. "Unemployment and Compensating Wage Differentials." Unpublished MS., University of Chicago, CMSBE, Chicago, 1979.

Arrow, K. *Essays in the Theory of Risk Bearing*. Amsterdam: North-Holland, 1970.

Ashenfelter, O. "Unemployment as Disequilibrium in a Model of Aggregate Labor Supply." *Econometrica* 48 (April 1980): 547–64.

Ashenfelter, O., and Heckman, J. "The Estimation of Income and Substitution Effects in a Model of Family Labor Supply." *Econometrica* 42 (January 1974): 73–85.

Azariadis, C. "Implicit Contracts and Underemployment Equilibria." *Journal of Political Economy* 83 (December 1975): 1183–1202.

Baily, M. N. "Wages and Employment under Uncertain Demand." *Review of Economic Studies* 41 (1974): 37–50.

Duncan, G., and Stafford, F. "Do Union Members Receive Compensating Wage Differentials?" *American Economic Review* 70 (June 1980): 355–71.

Feldstein, M. "Temporary Layoffs in the Theory of Unemployment." *Journal of Political Economy* 84 (October 1976): 937–57.

______. The Effects of Unemployment Insurance on Temporary Layoff Unemployment." *American Economic Review* 68 (December 1978): 834–46.

Heckman, J. J. "Sample Selection Bias as a Specification Error." *Econometrica* 47 (January 1979): 153–62.

Keeley, M. C., et al. "The Estimation of Labor Supply Models Using Experimental Data." *American Economic Review* 68 (December 1978a): 873–87.

______. "The Labor Supply Effects and Costs of Alternative Negative Income Tax Programs." *Journal of Human Resources* 13 (Winter 1978b): 3–36.

Lewis, H. G. "Interes del empleador en las horas de Trabajo del empleado" (Employer Interests in Employee Hours of Work). *Cuadernos de Economica* (Chile), 1969.

Lucas, R. E., and Rapping, L. A. "Real Wages, Employment and Inflation." *Journal of Political Economy* 77 (October 1969): 721–54.

Malinvaud, E. *The Theory of Unemployment Reconsidered.* N.Y.: John Wiley, 1977.

Mortensen, D. "Job Search, the Duration of Unemployment and the Phillips Curve." *American Economic Review* 60 (December 1970): 847–62.

Pratt, J. "Risk Aversion in the Small and in the Large." *Econometrica* 32 (January 1964): 122–36.

Rosen, S. Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition." *Journal of Political Economy* 82 (February 1974): 34–55.

______. "Substitution and Division of Labour." *Economica* 45 (August 1978): 235–50.

Survey Research Center. *A Panel Study of Income Dynamics: Waves I-V Procedures.* Ann Arbor: Institute for Social Research, 1972a.

______. *A Panel Study of Income Dynamics: Waves I-V Tape Codes.* Ann Arbor: Institute for Social Research, 1972b.

______. *A Panel Study of Income Dynamics: Wave VI Procedures and Tape Codes.* Ann Arbor: Institute for Social Research, 1973.

______. *A Panel Study of Income Dynamics: Wave VII Procedures and Tape Codes.* Ann Arbor: Institute for Social Research, 1974.

______. *A Panel Study of Income Dynamics: Wave VIII Procedures and Tape Codes.* Ann Arbor: Institute for Social Research, 1975.

______. *A Panel Study of Income Dynamics: Wave IX Procedures and Tape Codes.* Ann Arbor: Institute for Social Research, 1976.

Zellner, A. "Estimation of Functions of Population Means and Regression Coefficients Including Structural Coefficients: A Minimum Expected Loss (MELO) Approach." *Journal of Econometrics* 8 (October 1978): 127–58.

5 Structural and Reduced Form Approaches to Analyzing Unemployment Durations

Nicholas M. Kiefer and George R. Neumann

5.1 Introduction

Workers with low current earnings comprise two types of individuals: those whose personal characteristics lead to their being permanently in the low-wage state, and those who are, owing to some exogenous event, only transitorily in the low-wage state. This distinction is recognized implicitly in public policies designed to aid such workers. Workers who are viewed as "permanent" low wage earners are provided programs which attempt to alter their personal characteristics—e.g., manpower training programs. For those workers viewed as only transitorily in the low earning state, services provided tend to be short-term income maintenance, e.g., unemployment insurance following losses in jobs and Workmen's Compensation following debilitating work injuries. The distinction between permanent and transitory is not rigid, however, since not all workers recover from a transitory shock such as the loss of a high-wage job. Similarly, some workers with characteristics normally associated with permanent low wage earnings escape to the high-wage sector. The size of the pool of low wage at any time depends then upon the magnitudes of these inflows and outflows. Although economists cannot claim to understand fully how public programs affect all movements between the two states, a clearer picture is emerging on the effects of manpower training programs and the movement out of the low-earnings state.

Our understanding of the effect of public programs on the transition into the low-earnings state is much less precise, however, partially because we have only a limited knowledge of the adjustments individuals

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make to such events as job loss. Why is it, for example, that one individual will become reemployed in a short time with only minimal loss of earnings while another individual with a similar earnings history finds a new job only after a considerable period of time and then experiences a substantial decline in earnings? Is this merely an example of "bad luck," or does it indicate a systematic means whereby a transitory event leads some workers into permanent low-wage status? Although much has been written on the job search behavior of individuals, comparatively little empirical evidence exists to shed light on why some individuals succeed and others fail. Moreover, the evidence that does exist is generally of little use for exploring questions about the efficacy of alternative labor market programs. This latter problem arises because customary approaches of analyzing the outcome of the job search process—that is, the wage offer accepted and the length of time required to obtain it—produce, at best, a reduced form relationship which confounds differences in market opportunities with differences in personal characteristics. Consequently, the true effect of a particular program is difficult to determine. For the purposes of policy analysis, an identification of the underlying structural relationship is necessary if one desires to measure the effects of programs designed to affect the job search process.

In this paper we consider the effects of two alternative labor market programs designed to smooth the transition from the unemployed state: a modified version of regular unemployment insurance and a wage subsidy program. In the data used in this study, one of these programs—the modified unemployment insurance—actually operated, and we can therefore consider variations in policy parameters. The alternative wage subsidy program was not available to any individuals, but it has attracted some attention recently as a means of reducing unemployment. While no direct evidence—that is, of the experiences of treatment and control groups—is available, we show that knowledge of the *structural* parameters—but not the *reduced form* parameters—is sufficient to identify the effects of this type of program. In examining the effects of the different programs, we contrast the policy implications that flow from the reduced form estimates and the structural estimates. These differences provide a useful insight into the gains obtainable from a precise model specification.

5.2 Outcomes of the Job Search Process

Analysis of the effects of unemployment has focused on the length of time required to find employment, and the resulting wage obtained; in particular, the analysis has focused on measuring the effects of programs such as unemployment insurance (UI) on the outcome of the job search process. The theory motivating this analysis is given by the well-known papers by Mortensen (1970) and McCall (1970) on search behavior. To state this theory somewhat loosely, empirical studies proceed from the observation that anything which lowers the cost of search increases an individual's reservation wage and thereby leads to both longer durations of unemployment and higher wages upon reemployment.

Empirical efforts to measure the relationship between duration and wage change have taken two directions. The first approach, typified by Classen (1977) and Ehrenberg and Oaxaca (1976), treats the outcomes of the job search process as jointly determined and attempts to estimate a *reduced form* system. The specific model is:

$$D_i = X'_{1i}B + E_{1i} \tag{1a}$$

$$W_i = X_{2i}B_i + E_{2i} \tag{1b}$$

where D_i is the number of weeks of unemployment and W_i is reemployment wage. Parameters of the UI system, i.e., the replacement rate, are included in X_1 and X_2, and their coefficients are interpreted as the net effects of the UI system on the job search process.

An alternative approach has been taken in Kiefer and Neumann (1978 1979a b). In this approach the job search process is viewed as a selection problem following Heckman (1979). Individuals accept employment if and only if the market wage offer exceeds their reservation wage. Expected wages are then just a drawing from a truncated distribution, with the point of truncation depending upon the reservation wage, and the expected duration of unemployment is distributed geometrically about the inverse of the per period probability of finding an acceptable job offer.[1] A difficulty encountered in the approach is that reservation wages are not observable; they must be inferred from the observed choices of individuals. This problem, which motivated the use of a reduced form solution in other papers, can be solved in the following manner (see also Kiefer and Neumann 1979b).

Assume that the wage offer distribution facing the *i*th individual is:

$$\begin{aligned} &\ln w_{it}^o + X'_iB + f_i + \varepsilon_{it}^o \\ &\quad \varepsilon_{it}^o \sim \text{i.i.d. } N(0, \sigma_o^2)\ \forall t \end{aligned} \tag{2}$$

where X_i represents all measured characteristics of an individual (age, education, labor market characteristics, etc.), f_i represents all unmeasured characteristics, which are assumed known by the individual and potential employers, and ε_{it}^o is a random error term representing the "pure" amount of wage variability. The characterization in (2) implies that the wage offer distribution is stationary, an assumption which seems reasonable in light of the span of time covered by a typical spell of unemployment, and that observed wages have two sources of variation—systematic, but unmeasured, differences in "ability" f_i, and randomness in the wage offer process, represented by ε_{it}^o.

Facing (2), an optimal strategy is to select a reservation wage with the property that offers which match or exceed this critical value are accepted and those that fall short are rejected. The reservation wage can be shown to be of the form:

$$r_{it} = g[F(w_i^o), m, \theta, t] \tag{3}$$

where $F(w_i^o)$ is the distribution of wage offers, m is the direct cost of search, θ is the discount factor, and t represents the effect of state dependence—that is, reservation wages may systematically vary with the length of time searching. Using results from Kiefer and Neumann (1979a b), a first-order Taylor expansion of (3) can be shown to yield

$$r_{it} = k_i(X'_iB + f_i) + Z_i(t)\cdot\gamma \tag{3'}$$

where k_i is defined as

$$k_i = \frac{\int_{r_{it}}^{\infty} F(w_i^o)dw^o}{\int_{r_{it}}^{\infty} F(w_i^o)dw^o + \theta} = \frac{\alpha_i}{\alpha_i + \theta} = k_i(XB, Z\gamma, k_i)$$

Note that there is no stochastic element in (3′); individuals who search optimally in this model choose a strategy—a reservation price—which is not random, although it may vary over time as reflected in the time subscript on Z, i.e., in response to time-dependent factors which directly affect the costs of search.

Individuals accept employment if and only if the wage offer exceeds the reservation wage. Using (2) and (3′), the employment condition is that

$$s_i(t) = (1 - k_i)(X_i'B + f_i) - Z_i(t)\gamma > -\varepsilon_i^o \tag{4}$$

defining $s_i(t) = -[(1-k_i)(X_i'B + f_i) - Z_i(t)\gamma]/\sigma$. The probability of finding a job in any period α is, for a given individual,

$$\alpha[s_i(t)\,|\,f_i] = \Pr(w_i^o > w_i^r\,|\,f_i) = 1 - \Phi[s_i(t)\,|\,f_i] \tag{5}$$

where Φ is the standard normal distribution function. The statement in (5) is the probability of an individual's finding a job in period t, conditional on his unmeasured ability f_i. Although by definition we do not have measures of f_i, an implication of the optional choice of a reservation wage is that randomness in wage offers should be independent of f_i. Hence the unconditional probability of finding an acceptable job offer is

$$\alpha[s_i(t)] = \int_{-\infty}^{\infty}[1 - \Phi(s_i\,|\,f_i)]d\Phi\left(\frac{f_i}{\sigma_F}\right) \tag{6}$$

Using (6) and results from conditional normal theory, the probability of observing a particular outcome—that is, a wage w_i^o, and a length of unemployment D_i—is given by

$$\text{(7)} \qquad \Pr(w_i^o, D_i) = \int_{-\infty}^{\infty} \left(\{\Pi\ \Phi[s_i(t) \mid f_i]\} \cdot \frac{1}{\sigma_o} \phi(\varepsilon_i) \cdot [1 - \Phi(s_i \mid f_i]\right) d\Phi\left(\frac{f_i}{\sigma_F}\right)$$

as t goes from 1 to $D_i - 1$ and $d\phi(f_i/\sigma_F)$ goes from $-\infty$ to ∞. Subject to identification criteria discussed in Kiefer and Neumann (1979a b), all parameters in equation 7 can be estimated by maximum likelihood methods.[2] In particular one can identify B, γ, σ_o^2 (the pure variation in wage offers), and σ_F^2 (the variation in unmeasured ability).

The issues which arise in estimating the model described above are discussed at length elsewhere (see Kiefer and Neumann 1979b). For the present purposes it is sufficient to note that two structural equations relating unemployment and reemployment earnings are embedded in (7). The expected length of search for a randomly chosen individual is given by:

$$\text{(8)} \qquad E(D_i) = \sum_{j=1}^{\infty} \int_{-\infty}^{\infty} \left(\prod_{\ell=1}^{j-1} \Phi[s_i(\ell) \mid f_i] \right) [1 - \Phi(s_i(j) \mid f_i] \cdot j\ d\Phi\left(\frac{f_i}{\sigma}\right)$$

The expected reemployment wage is somewhat more cumbersome to derive. Conditional on f_i, and conditional on the length of search being D_i, expected reemployment earnings are:

$$\text{(9)} \qquad E(w_i^o \mid f_i, D_i) = X_o'B + f_i + \sigma_o \lambda[s(D) \mid f_i]$$

where

$$\gamma[s(D) \mid f_i] = \frac{\phi[s_i(D) \mid f_i]}{1 - \Phi[s_i(D) \mid f_i]}$$

If the reservation wage were constant, i.e., s did not vary with D, then unconditional expected earnings would be given by

$$\text{(10)} \qquad E(w_i^o) = \int_{-\infty}^{\infty} E(w_i \mid f_i, D) d\Phi\left(\frac{f_i}{\sigma}\right) = X_i'B + \sigma_o \int_{-\infty}^{\infty} \lambda(s_i \mid f_i) d\Phi\left(\frac{f_i}{\sigma}\right)$$

When reservation wages vary with search time, the second term on the right-hand side on (10) must be modified to allow for differences in the probability of receiving an acceptable offer in a given period. Define the probability that an acceptable offer is received in period j as:

$$(11) \qquad g_i(j) = \left(\prod_{\ell=1}^{j-1} \Phi[s(\ell) \mid f_i]\right) \cdot 1 - \Phi[s(j) \mid f_i]$$

The unconditional expected reemployment wage is then:

$$(12) \qquad E(w_i^o) = \int_{-\infty}^{\infty} \left(\sum_{j=1}^{\infty} E(w_i^o \mid f_i,\ D=j) \cdot g_i(j)\right) d\Phi\left(\frac{f_i}{\sigma_o}\right)$$

$$- X_i'B + \sigma_o \cdot \int_{-\infty}^{\infty} \left(\sum_{j=1}^{\infty} \lambda[s_i(j) \mid f_i] \cdot g_i(j)\right) d\Phi\left(\frac{f_i}{\sigma_o}\right)$$

Equations 8 and 12 can be thought of as the structural analogues to what we have termed the reduced form solutions of (1a) and (2b). In view of the differences between the reduced form and structural approaches it is useful to examine the merits of each. Two issues are of particular importance: interpreting changes in policy variables such as UI benefits, and drawing inferences from incomplete samples (see Johnson and Kotz 1972; Heckman, in press).

The reduced form approach has one particular advantage—it is simple and cheap to estimate. If reservation wages are constant, the estimated coefficients have a potential interpretation as the coefficients of a Taylor expansion of the inverse of (6) for the duration equation [i.e., $E(D) = 1/\alpha(s_i)$], and as

$$B + \sigma_o \int_{-\infty}^{\infty} \frac{\partial\gamma(\cdot)}{\partial(\cdot)} \frac{\partial(\cdot)}{\partial r} \frac{\partial n}{\partial X_i} \, d\Phi\left(\frac{f_i}{\sigma}\right)$$

for the earnings equation. In this case, if both forms of the job search model were estimated on a complete sample, the only difference that should arise would be due to the inherent nonlinearity of the structural duration equation. If reservation wages vary over time as well as across individuals, then the correspondence between the two approaches is less obvious. Policies which affect the duration of unemployment also affect the distribution of accepted wages since the point of truncation varies with duration.

The use of a reduced form approach also results in problems of interpretation when certain types of policy simulations are attempted. For example, if a wage subsidy of, say, ten percent were given to all individuals in the sample, it would affect both duration and reemployment earnings, although in opposite ways. In the absence of a controlled experiment—where individuals were randomly assigned to the group receiving the subsidy—it is difficult to see how one could simulate this effect using a reduced form model. The problem is one of identification:

the moments of the wage distribution do not enter explicitly into the reduced form approach. If reservation wages are constant, this problem may not be serious because of the potential interpretation of the reduced form coefficients noted above. In the more general case, however, it is not possible to infer the results of such an experiment from the reduced form estimates.

Perhaps the greatest difference between the two approaches arises when information is available only for an incomplete sample. For example, it is frequently the case that a "follow-up" survey is performed after some event has occurred. At the time of the survey some individuals will have completed their job search, but some will not. Those who have not found employment will tend to have low expected market earnings, relative to their reservation wage—hence the long period of unemployment. Since neither of the dependent variables is observed, the observations are usually excluded from the analysis.[3] For well-known reasons this is likely to result in biased estimates. Apart from the question of bias, there is the question of interpreting the results of any simulation exercise since the composition of an incomplete sample is not likely to be invariant under changes in policy. Consider, for example, the effect of a shift in the mean of the wage offer distribution. Search theory implies that the expected wage should increase, and expected duration decrease, for all individuals. In an incomplete sample, the effect of such a policy would be that some individuals who previously had not found employment would become employed and hence would be included in the sample. If these individuals on average had higher durations of unemployment and lower expected earnings, then *observed* average wages would fall and duration increase, even in a carefully controlled experiment.

The importance of this effect will depend upon the location of reservation wages along the distribution of wage offers. If reservation wages are high, relative to the mean of the wage offer distribution, and if the distribution of offers has small variance, even a small shift in the mean may produce a significant change in unemployment patterns.

In noting these differences, we have only pointed out the potential problems which may exist; the severity of these problems—that is, the extent to which they lead to different policy implications—is ultimately an empirical matter. In the following section, we examine the simulated responses of a group of individuals to two plans which affect their unemployment activities.

5.3 Simulating Job Search Behavior: The Effects of A Wage Subsidy Plan

In this section we apply the models discussed above to a sample of unemployed male workers. This particular sample was generated from a survey of trade-displaced workers conducted by the Institute for Research on Human Resources of the Pennsylvania State University. A complete description of the data source is contained in Neumann (1978). Several features make this group particularly appropriate for discussions about low-wage workers. The sample is constructed solely of individuals who were permanently separated from employment—in most cases because the entire plant shut down. Thus we observe only job search behavior and do not have to be concerned with responses to anticipated, temporary layoffs. Moreover, the nature of the shock conforms to the idea of an exogenous shock to which some individuals adjust reasonably well, and others adjust only with great difficulty. Although many of these individuals would not have been considered low-wage workers prior to displacement, the average loss in weekly earnings upon reemployment was over twenty-five percent: consequently, most would be considered low-wage earners afterward. Summary statistics on this sample are contained in table 5.1.

Estimates of the reduced form equations for duration and reemployment earnings are presented in table 5.2, and the structural estimates of reemployment earnings (wage offers) and reservation wages are contained in table 5.3. Although we will not dwell on the precision of the estimates, we do note that the explanatory power of the OLS regression of unemployment duration is exceedingly small; this appears to be a common finding (see, e.g., Ehrenberg and Oaxaca 1976; Classen 1977).

Both approaches indicate an effect of UI benefits on the outcome of the job search process. The reduced form estimates imply that a ten percent increase in the replacement rate—equivalent here to an average increase of \$14.9 per week in UI benefits—would lead to an increase in duration of about one-half week ($.0314 \times 14.9$), and an increase in unemployment earnings of 0.60 percent. The effects of increased UI benefits are apparent in column 2, but the numerical values of the increases in duration and reemployment earnings depend upon the position of the reservation

Table 5.1 **Sample Characteristics of Male Workers**

	Mean	Maximum	Minimum
Education (years)	10.2	21.0	0.0
No. of dependents	1.7	9.0	0.0
Percent married	83.5	—	—
Percent union members	70.4	—	—
Local unemployment rate at layoff (%)	5.30	9.00	2.20
Age	47.8	75.0	19.0
Unemployment benefits per week (\$1967)	62.7	117.11	0.0
Maximum benefit period (weeks)	41.5	65.6	0.0
Previous weekly earnings (\$1967)	149.0	457.0	19.20

Table 5.2 Reduced Form Estimates of Duration and Reemployment Wage Equation

	Duration (1)	Reemployment Earnings (2)
Constant	18.1566	1.839
	(2.17)	(3.16)
Education	0.0161	0.0088
	(0.96)	(2.41)
Dependents	0.0261	0.0617
	(0.41)	(0.14)
Tenure	0.0040	−0.0069
	(1.06)	(1.92)
Marital status	0.0001	0.1139
	(0.00)	(0.60)
Unemployment rate	2.1164	−0.0461
	(1.97)	(1.27)
Age	+0.0441	0.0210
	(1.40)	(1.21)
Age^2	−0.0003	−0.0002
	(0.27)	(0.06)
Ed · Age	+0.0143	−0.0011
	(0.20)	(1.61)
UI benefits	0.0314	0.0004
	(1.71)	(1.30)
Maximum duration	0.0214	−0.0001
	(1.40)	(0.01)
$\ell n(W_{t-1})$	−0.3118	0.5406
	(1.11)	(7.24)
R^2	.1331	.2480
F	1.478	9.012

wage in the wage offer distribution. We calculate these effects in the simulation reported below.

Before examining the simulation results it is useful to consider one feature of the job search process. Both casual empirical evidence and some previous studies (e.g., Neumann 1978) suggest that losses due to unemployment are greatest for the long-term unemployed. Although a higher reservation wage leads to higher expected reemployment and a greater length of unemployment for any individual ex ante, when one observes the outcomes of the job search process ex post, this investment

Table 5.3 **Structural Estimates of the Job Search Process**

	Earnings Function (1)	Reservation Wage Function (2)
Constant	2.8263	1.9713
	(6.24)	(3.47)
Education	0.0361	0.0101
	(1.87)	(1.27)
Dependents	—	−0.0068
		(0.47)
Tenure	−0.0078	—
	(3.68)	
Marital status	—	−0.0824
		(3.68)
Unemployment rate	0.0197	0.0161
	(1.68)	(2.89)
Age	0.0194	−0.0127
	(1.86)	(3.46)
Age^2	−0.0001	0.0001
	(0.61)	(0.84)
Ed·Age	−0.0008	−0.0003
	(1.87)	(1.71)
Unemployment benefits	—	0.0016
		(2.43)
Maximum duration	—	0.0004
		(0.59)
$\ell n W_{t-1}$	0.2574	—
	(4.57)	
F_i	—	−0.0014
		(0.91)
t	—	−0.0023
		(2.01)
$\sigma^2_{W^o}$	0.0283	
	(2.62)	
σ^2_F	0.2493	
	(12.41)	
$\ell n\ \ell\ell$	−1,794.83	

Note: t-statistics in parentheses.

aspect is swamped by variations in individual characteristics and by random errors in the process. In the present context this phenomenon is likely to be concentrated among the group of workers who had not found employment by the survey date. Since their behavior is of particular interest in any discussion of low-income workers we present simulation results separately for this group.

The simulated effects of changing UI benefits in steps of five percent on duration of unemployment and the percentage change in reemployment earnings are presented in table 5.4. Panels A and B contain the estimates from the reduced form model (equations 1a and 1b) for the total sample and for those workers who remained unemployed for at least sixty-five weeks; panels C and D contain the equivalent estimates for the structural model (equations 8 and 12). The estimates in table 5.4 show two pronounced patterns. Looking across each panel, we see that, for this sample at least, changes in UI benefit levels would have almost negligible effects. Increasing UI benefits by twenty percent—which for this sample is equivalent to raising the average replacement rate by 8.4 percentage points (from 42.1 percent to 50.5 percent)—would raise reemployment earnings by only about .5 percent and increase the duration of unemployment by about one-half week. These are quite modest effects when one considers that the average reemployed worker in this sample had a decline in real weekly earnings of 26.7 percent and spent 39.1 weeks unemployed. It is interesting to note that although estimates of the precise effect of changing UI benefits would differ depending upon whether one used the reduced form or structural model, the conclusions to be drawn from the evidence would not.

Looking down the columns of table 5.4, we observe a somewhat different picture of the differences between the two approaches to modeling the job search process. Comparison of panels A and B would seem to indicate that there is little difference between those who had not become employed within 65 weeks and those who had; panels C and D indicate the contrary. The expected duration of unemployment was estimated to be 34.7 weeks for those who became employed within 65 weeks, and 47.2 weeks for those who had not become employed by 65 weeks. This amounts to about a seven-week difference in expected duration of unemployment between the two groups. In one sense, this difference between the two models can be considered a contrived one, since the structural model takes into account information on the characteristics and, partially, the job search outcomes, of the group of workers who had not found jobs within 65 weeks.[4] But this is precisely the purpose of a structural model, and the differences observed in table 5.4 represent the basis for using such an approach to design policies to smooth labor market transitions. Under the reduced form approach, the similarity of the estimated duration and wage changes would lead one to conclude that

Table 5.4 **Structural and Reduced Form Simulations of the Effect of Alternative Levels of UI Benefits**

	% Δ in UI Benefits				
	0.0	5.0	10.0	15.0	20.0
	Reduced Form Estimates A. Total Sample				
Duration (weeks)	39.31	39.41	39.51	39.61	39.72
% Δ in earnings	0.0	0.13	0.25	0.38	0.50
	B. Unemployed after 65 weeks				
Duration (weeks)	39.62	39.73	39.83	39.93	40.03
% Δ in earnings	0.0	0.13	.025	0.37	0.49
	Structural Estimates C. Total sample				
Duration (weeks)	43.10	43.41	43.67	43.85	43.91
% Δ in earnings	0.0	0.17	0.29	0.46	0.54
	D. Unemployed after 65 weeks				
Duration (weeks)	47.21	47.36	47.50	47.61	47.71
% Δ in earnings	0.0	0.11	0.18	0.25	0.31

the two groups are essentially the same; hence it must be random influences—luck—which determine whom the labor market assigns to each group. The structural approach, on the other hand, implies that there are real differences between the two groups and thus, at least in principle, allows the possibility of predicting in advance what types of individuals are likely to be most affected by unexpected job loss.

The results of this simulation raise strong doubts about the ability of what is essentially an income maintenance program to have a significant impact on the reemployment experience of displaced workers. Although the sample used is unique, and certainly not representative of all unemployed workers, our results, both the reduced form and structural versions, are not significantly at odds with the findings of others which are based solely on a reduced form approach. While it is difficult to generalize from a sample of one, there is at least the suggestion that returns from more precise modeling of the job search process may be important for policy purposes.

Although predicting which types of individuals will be most adversely affected by job termination is one possible gain to a structural approach, a more important gain is likely to be in terms of the number of difference policy options which can be considered. As an example, we consider the option of a wage subsidy program. The basic idea of a wage subsidy is to shift the distribution of wage offers facing individuals, thereby making employment more likely. In the reduced form approach there is no obvious way to incorporate such effects, except possibly through a controlled experiment. A structural approach allows for a direct interpretation, however, since the shift in the wage offer distribution affects an individual's expected earnings both directly—i.e., through $X_i'B$—and indirectly through its effects on reservation wages.

In table 5.5 we present the results of a simulation exercise with varying amounts of wage subsidy. Because these simulations, as in the case of the UI subsidy, are partial equilibrium in nature, the results are sensitive to the assumed stability of the wage offer distribution. In the present case, this amounts to assuming that a wage subsidy program will not affect the distribution of wage offers part from the mean shift, i.e., no "extra" effects due to a substitution of labor for capital. For small programs this assumption seems tenable.

The issue also arises of how accurately this shift in the distribution is perceived by individuals. If it is fully perceived, then reservation wages rise by a fraction $\alpha/\alpha + \theta$ of the increase in the mean. This increase in reservation wages leads to lengthier search, and, consequently, the effect on duration of unemployment is lessened. Since some wage subsidy plans (e.g., jobs credit) work in a manner that may not be obvious to individuals, we present estimates of the effect on duration assuming full reservation wage change (panels A and B), and no reservation wage change (panel C).

In contrast to a UI subsidy, a direct wage subsidy appears to have quite significant effects on the job search process. From panels A and B we observe that a twenty percent wage subsidy would lead to an increase in reemployment earnings of about nineteen percent, and a reduction of unemployment duration of about a week, if the shift in the mean is

Table 5.5 **Structural Simulations of the Effect of a Wage Subsidy Program**

	% Δ in Mean Wage Offer				
	0.0	5.0	10.0	15.0	20.0
	A. Total Sample				
Duration (weeks)	43.10	42.87	42.51	42.23	42.06
% Δ in earnings	0.0	4.91	9.84	14.72	19.6
	B. Unemployed over 65 weeks				
Duration (weeks)	47.21	47.03	46.74	46.39	46.12
%Δ in earnings	0.0	4.87	9.78	14.68	19.2
	C. Duration of Unemployment with Incomplete Knowledge (weeks)				
Total sample	43.10	41.64	40.02	38.75	37.29
Unemployed over 65 weeks	47.21	45.88	44.16	42.82	41.28

completely perceived. The effect of the change in reservation wages can be seen clearly in panel C: if reservation wages did not adjust, expected unemployment duration would decrease by six weeks instead of one.

5.4 Conclusion

This paper has focused on two points—the inferences which can be obtained from structural versus reduced form analysis of the outcome of the job search process, and the effects of two subsidy programs on the job search process. In regard to the former topic, it is clear that a structural model permits a wider range of possible questions. In particular, it is possible to consider, ex ante, what the likely experience of a given cohort of job searchers will be, and, in principle, to tailor different types of programs to ease their labor market transitions.

The comparison of a UI subsidy with a wage subsidy revealed significant differences. Higher levels of UI payments led, as expected, to both longer durations of unemployment and higher reemployment earnings. Both effects were quite small, however, and, at least for low-wage workers similar to the individuals in this sample, there is little reason to believe that programs which emphasize income maintenance are likely to have much impact on the types of jobs obtained. By contrast, a wage subsidy program appears to have a significant effect on reemployment earnings, and also to lead to a moderate decline in duration. This is a one-blade-of-the-scissors result of course, and it is subject to criticism on those grounds. Nonetheless, for relatively small programs, the possibilities appear to be fruitful.

Notes

1. This result holds only for the case of constant reservation wages. The correct distribution of durations for the general use is given in equation (7) below.
2. The identification criteria amount to the following: some variable(s) must affect wage offers but must not directly affect reservation wages. Indirect effects—e.g., through the moments of the wage offer function—are permissible, indeed necessary.
3. There are other reasons why truncation could occur. Using state UI records on compensated unemployment results in a truncation of those with very short durations—less than the waiting period—and those with long durations—those whose unemployment exceeds the maximum duration period.
4. The estimates in the reduced form approach for the sample of workers not employed in sixty-five weeks are constructed simply by using the observed characteristics of the individual and the coefficients estimated from the sample of employed. No attempt is made to adjust the constant term such that the expected value of, say, duration reflects the obvious fact that the observed period of unemployment was greater than 65 weeks.

References

Classen, K. "The Effect of Unemployment Insurance on the Duration of Unemployment and Subsequent Earnings." *Industrial and Labor Relations Review* 30, no. 4 (July 1977): 438–44.

Ehrenberg, R., and Oaxaca, R., "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Gain." *American Economic Review* 66 (December 1976): 754–66.

Heckman, J. "Sample Selection Bias as a Specification Error." *Econometrica* 47 (January 1979): 153–62.

Johnson, N., and Kotz, S. *Distributions in Statistics*. Vol. 4: *Continuous Multivariate Distributions*. N.Y.: John Wiley, 1972.

Kiefer, N., and Neumann, G. "Estimation of Wage Offer Distributions and Reservation Wages," in S. Lippman and J. McCall, eds., *Studies in the Economics of Search*. Amsterdam: North-Holland, 1978, pp. 171–89.

———. "An Empirical Job Search Model with a Test of the Constant Reservation Wage Hypothesis." *Journal of Political Economy* 87, no. 1 (February 1979a).

———. "Individual Effects in a Nonlinear Model: Explicit Treatment of Heterogeneity in the Empirical Job-Search Model." Mimeo., February 1979.

McCall, J. "Economics of Information and Job Search." *Quarterly Journal of Economics*, February 1970, pp. 113–26.

Mortensen, D. "Job Search, the Duration of Unemployment, and the Phillips Curve," *American Economic Review* 60 (December 1970): pp. 847–62.

Neumann, G. "The Labor Market Adjustments of Trade Displaced Workers: The Evidence from the Trade Adjustment Assistance Program," in R. Ehrenberg, ed., *Research in Labor Economics*, pp. 353–81. JAI Press, 1978.

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