NBER WORKING PAPER SERIES

COMPENSATING WAGE DIFFERENTIALS AND THE DURATION OF WAGE LOSS

Daniel S. Hamermesh

John R. Wolfe

Working Paper No. 1887

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
April 1986

Helpful comments were provided by Jeff Biddle and Carl Davidson and by participants in seminars at several universities. The research reported here is part of the NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

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ABSTRACT

Several reasons are offered why workers will receive larger compensating wage differentials for increases in the duration of wage losses than for increases in the probability of loss that produce the same expected loss. A formal model of occupational choice is developed that shows the extent to which the compensation for increased duration exceeds that for increased risk.

Using Panel Study of Income Dynamics data linked to industry data on injuries and unemployment, we find: 1) Nearly all the compensating wage differential for losses due to workplace injuries is compensation for increases in the duration of loss; 2) Similarly, nearly all the compensation for losses due to cyclical unemployment is compensation for increases in duration, especially for increases in duration beyond the 26 weeks of unemployment that are usually compensated by unemployment insurance. The compensating differentials for risk of injury are larger for union than for nonunion workers, while those for cyclical unemployment are smaller for union workers.

Daniel S. Hamermesh
Department of Economics
Michigan State University
East Lansing, MI 48824
(517) 355-5238

John R. Wolfe
Department of Economics
Michigan State University
East Lansing, MI 48824
(517) 355-1863

The concept of compensating wage differentials has been a hardy device for generating fruitful hypotheses about wage structure. Clearly, not all hypotheses growing out of the concept will be supported by the evidence, nor will all of them apply at all places and at all times. Rees (1975)

I. Introduction

A substantial and rapidly growing body of research has examined the relationship between wage levels and the likelihood of wage loss. With its roots in Adam Smith, (1937, Book I, Chapter 10), this literature has studied the effect of the risk of work-related fatal accidents (Thaler-Rosen, 1975; Smith, 1979); work-related nonfatal accidents (Viscusi, 1979; Olson, 1981); and the risk of unemployment (Abowd-Ashenfelter, 1981; Topel, 1984; Li, 1986). The entire genre of research examines how wages among otherwise identical individuals differ as their expected wage losses vary.

Though the literature clearly stems from Adam Smith, the "hardy device" in the Wealth of Nations has not, we believe, been fully exploited to generate all the "fruitful hypotheses" in this area that it might. Any expected wage loss is made up of two components: The incidence of the loss--the probability that the loss will occur; and the duration of loss conditional on its occurrence. Two otherwise identical workers can face the same expected loss, yet face sharply differing incidence and duration of loss. For several reasons we should not expect these two workers to receive identical compensating differentials for the same expected loss. A worker's preferences will not be symmetric in frequency and duration, unless the worker is risk neutral: A doubling of duration will provide greater disutility to a risk-averse worker than will a doubling of frequency, because such a change will broaden the distribution of possible wage losses. Insofar as borrowing to finance consumption is difficult, and the worker is illiquid, this difference will be especially pronounced. Similarly, as Hurd (1980) and Layard (1982) argue, upward-sloping labor supply curves guarantee that the expected utility loss arising from enforced leisure of a given expected length is greater if the loss is a long-duration, low-probability event than if it is a brief, high-probability occurrence.

The strongest indication in the empirical literature on wage losses that duration plays an especially important role is the evidence that compensation for risk of death is much greater than that for finite losses (Smith, 1979). Such a strong aversion to the risk of death is the limiting case of the phenomenon that we propose: An aversion to the risk of large losses, holding the expected loss constant.[1]

In this study we develop a model that derives the effects of variations in the incidence and duration of wage loss on workers' expected lifetime utilities. The relationships imply a locus of equilibrium combinations of wages and expected wage losses that depends on the duration of loss conditional upon a loss occurring. The model is then compared to the standard model that ignores the distinction between incidence and duration and also to less structured estimating equations. Data from the Panel Study of Income Dynamics are used along with published data by industry on the incidence and duration of various types of loss. Our purpose is not to suggest that one particular mechanism produces unequal wage responses to incidence and duration of loss; rather, it is to suggest one such model, then to examine whether in fact the responses are unequal.

II. The Importance of Duration

The expected duration of wage loss from an event of given severity is the product of the duration of wage loss and the frequency of the event. Let jobs vary according to the risk of wage loss, parameterized as follows: Let the per-period probability of a loss occurring be β, and the duration of the wage loss be γ periods. Assume also that the wage w is a differentiable function of β and γ, and that wages are replaced in proportion $[1-\alpha]$ by social insurance during periods of wage loss. Let the worker's career be of fixed length T. Given γ for a job choice, we can redefine the career to consist of T/γ periods of length γ; during each such period the probability of wage loss is $\beta\gamma$. For each period t, define a random variable L(t) such that:

$$(1) \qquad L(t) = \begin{cases} 0 \text{ if no wage loss occurs} \\ 1 \text{ otherwise .} \end{cases}$$

L(t) is binomially distributed, with mean $\beta\gamma$ and variance $\beta\gamma[1-\beta\gamma]$.

Suppose that workers are risk averse with respect to lifetime consumption. In particular, assume the following utility function:

$$(2) \qquad U(c_1, c_2, \ldots, c_{\frac{T}{\gamma}}) = \frac{1}{1-\delta} \sum_{t=1}^{\frac{T}{\gamma}} \left[\frac{c(t)}{[1-\rho]^{\gamma t}} \right]^{1-\delta} ,$$

where ρ is the rate of time preference and δ is the degree of relative risk aversion. We avoid the unnecessary complications of borrowing and lending by assuming a constant marginal rate of substitution across periods and a rate of time preference that is equal to the interest rate. A simple consumption plan in which each period's consumption equals current income then maximizes utility, since the worker is indifferent between all consumption plans that exhaust lifetime income. We are thus free to focus on workers' choices with respect to the distribution of possible discounted lifetime incomes.

Given these assumptions, each period's consumption is:

(3) $$c(t) = \gamma[w[1-L(t)] + [1-\alpha]wL(t)]$$
$$= \gamma w[1-\alpha L(t)] \ .$$

The distribution of c(t) is therefore:

(4) $$c(t) \sim B(\gamma w[1-\alpha\beta\gamma],\ \gamma^2 w^2 \alpha^2 \beta\gamma[1-\beta\gamma]) \ .$$

The worker's problem is then to maximize expected lifetime utility by choosing job characteristics β and γ which in turn determine the wage and the probability of wage loss.

If we assume the L(t) are independently and identically distributed, then the mean and variance of lifetime consumption are easily determined. For convenience, define the asymptotically normal variable:

(5) $$C \equiv \sum_{t=1}^{\frac{T}{\gamma}} \frac{c(t)}{(1-\rho)^{\gamma t}} \sim n\left(\sum_{t=1}^{\frac{T}{\gamma}} \frac{\gamma w[1-\alpha\beta\gamma]}{[1+\rho]^{\gamma t}},\ \sum_{t=1}^{\frac{T}{\gamma}} \frac{\gamma^2 w^2 \alpha^2 \beta\gamma[1-\beta\gamma]}{[1+\rho]^{2\gamma t}}\right) \ .$$

For ρ small and T/γ large, the mean and variance can be approximated by:

(6) $$C \sim n\left(\frac{w[1-\alpha\beta\gamma]}{\rho},\ \frac{\gamma w^2 \alpha^2 \beta\gamma[1-\beta\gamma]}{2\rho}\right) \ .$$

Note that mean lifetime consumption is symmetric in β and γ, but that the variance of C is not: Potential losses of longer duration make lifetime income and consumption more uncertain, even if frequency is lowered by an equal proportion.

Each worker chooses among job characteristics β and γ in order to maximize:

(7) $$E(U(C)) = \int_{-\infty}^{+\infty} \frac{C^{1-\delta}}{1-\delta} f(C)dC,$$

subject to (6), where f is the density of C. The term $C^{1-\delta}/[1-\delta]$ can be approximated using a second-order Taylor-series expansion around $C=\mu$:

(8) $$\frac{C^{1-\delta}}{1-\delta} \approx \frac{\mu^{1-\delta}}{1-\delta} + \mu^{-\delta}[C-\mu] - \delta\mu^{-\delta-1}\frac{[C-\mu]^2}{2} .$$

Thus:

(9) $$E(U(C)) \approx \frac{\mu^{1-\delta}}{[1-\delta]} - \frac{\delta\mu^{-\delta-1}}{2}\sigma^2 = \frac{\mu^{1-\delta}}{[1-\delta]}[1-\frac{\delta[1-\delta]\sigma^2}{2\mu^2}] ,$$

which increases with the mean of lifetime consumption but decreases with its variance σ^2.

Substituting for μ and σ^2, we obtain:

(10) $$E(U) = [\frac{w[1-\alpha\beta\gamma]}{\rho}]^{1-\delta}[\frac{1}{1-\delta}][1 - \frac{\delta[1-\delta]}{2}\frac{\gamma\rho\alpha^2\beta\gamma[1-\beta\gamma]}{2[1-\alpha\beta\gamma]^2}] .$$

Note that substitution for σ^2 in the final term renders expected utility asymmetric in β and γ. Maximizing E(U) with respect to β and γ is equivalent to maximizing:

(11) $$\ln(1-\delta)E(U) = [1-\delta]\,[\ln w + \ln[1-\alpha\beta\gamma] - \ln\rho] + \ln(1-(\frac{1}{4})[\delta[1-\delta][\rho\alpha^2\beta\gamma^2]\,[1-\beta\gamma][1+\alpha\beta\gamma + [\alpha\beta\gamma]^2 + \ldots]^2)$$

If we assume $\beta\gamma$ to be small, ignore second- and higher-order terms in $\beta\gamma$, and approximate $\ln[1+x]$ by x for small values of x, we can write (11) as:

(12) $$\ln(1-\delta)E(U) = [1-\delta][\ln w-\alpha\beta\gamma-\ln\rho] - \frac{1}{4}\delta[1-\delta]\rho\alpha^2\beta\gamma^2$$

The first-order conditions for maximization with respect to β and γ are:

(13) $$\frac{\partial \ln w}{\partial \beta} = \alpha\gamma + \frac{\delta\rho\alpha^2\gamma^2}{4} ,$$

and:

(14) $$\frac{\partial \ln w}{\partial \gamma} = \alpha\beta + \frac{\delta\rho\alpha^2\beta\gamma}{2} .$$

Equations (13) and (14) describe the point chosen by the worker from the available frontier $w(\beta,\gamma)$ of job opportunities.[2] The following function, which satisfies both (13) and (14), therefore describes the wage frontier in the neighborhood of any chosen job:

(15) $$\ln w = \ln w_o + \alpha\beta\gamma + \frac{\delta\rho\alpha^2\beta\gamma^2}{4} ,$$

where w_o is the worker's wage in an occupation in which $\beta=\gamma=0$.[3] Wages should be a log-linear function of $\beta\gamma$, the expected fraction of earnings lost and not replaced, and of a risk-aversion term in which duration plays a more important role than does incidence.

A simplified graphical exposition can illustrate the main points of our argument. Following Abowd-Ashenfelter (1981), let $V(W)$ be the indirect utility of wage W per period. Arbitrarily assume that $\beta=1$ and $\gamma=\bar{\gamma}$. Then if the full-time per-period wage is $W*$, the per-period indirect utility is $V(W*[1-\bar{\gamma}])$, point B in Figure 1. Assume there is another industry such that $\gamma=2\bar{\gamma}$ and $\beta=1/2$, so that the expected wage loss remains constant across the two industries at $\bar{\gamma}W*$. Then the indirect utility in the second case is at point A in Figure 1. To attract workers to this second industry a wage sufficiently higher than $W*$ must be paid. The required wage is $W** > W*$, such that the expected indirect utility, the average of the indirect utilities obtained when no loss is incurred and when the loss is of duration $2\bar{\gamma}$, is equal to that

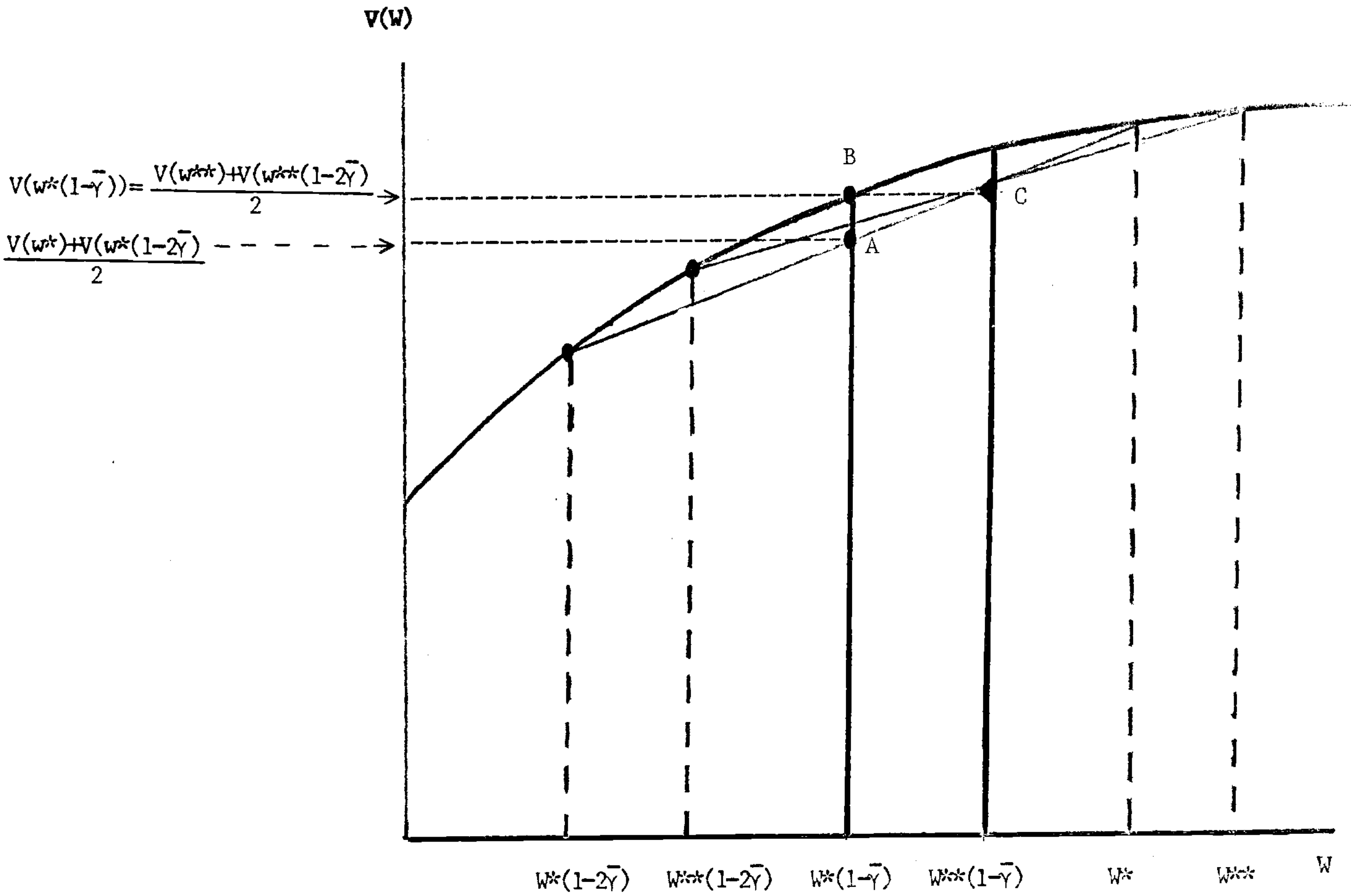

Figure 1.

The Need For Extra Compensation For Duration of Loss attained with a wage of $W*$ and a certain loss of shorter duration, $\bar{Y}$.

Attempts to estimate the compensating differential due to possible wage loss typically are of the same form as (15), less the final term. By forcing incidence and duration to have symmetric effects, this specification subjects the perceived costs of wage loss to measurement error, and results in a tendency to underestimate the effect of these perceived costs on wages.

III. An Application to Occupational Injuries

In this section we apply the model to data on the incidence and duration of workplace injuries. One study--Dorsey (1983)--did include measures of both the frequency and severity of injury in equations "explaining" wage differentials, but gave no reason for doing so and paid no attention to their separate effects. This particular application is thus the first test of the notion that risk and duration of loss will produce unequal compensating wage differentials for workplace injuries.

Equation (15) must be modified for estimation. An empirical version of the model in (15) is:

$$\ln W = \alpha_1 LW + \alpha_2 LW \cdot DUR + \beta X + \varepsilon \ . \tag{16}$$

where DUR is the duration of loss; LW=DUR·INC is the expected loss, the product of incidence and duration; X is a vector of other variables, and ϵ is a disturbance term. The difficulty with this estimating equation is that it specifies the separate effects of duration and incidence quite restrictively. Accordingly, we also estimate:

$$\ln W = \alpha_1' \, [\ln DUR + \alpha_2' \ln INC] + \beta X + \varepsilon \ . \tag{17}$$

Equations (16) and (17) are estimated by ordinary least squares. The data on which the estimation is based describe heads of households in the Panel Study of Income Dynamics who were between the ages of 22 and 65 in 1981.

This set of data provides no information on workers' assessments of risks on the job. Instead, we link the PSID data to published injury data, with the link based on industry affiliation in 1981, since that is the only year for which detailed industry data are provided.[4] The equations are estimated separately for 1980 and 1981.

The three-digit code identifying the industry to which the worker's employer belonged was used to link the record for the worker to Bureau of Labor Statistics data on workplace injuries.[5] While the correspondence between the two codes was not perfect, departures from a perfect match disqualified relatively few observations. This problem and the lack of complete information on all the variables required for the vector X resulted in a sample of 1689 household heads for 1981, and 1497 for 1980. Insofar as workers report their industry affiliation incorrectly, estimates of the compensating differentials will be biased toward zero, with a bias that Mellow-Sider (1983) show can be fairly large.[6]

The means of the BLS injury data in this sample of individuals are shown in Table 1. The figures on incidence and expected days lost (LW) are per 100 worker-years. The incidence data imply that roughly five percent of the workforce experiences at least one day of lost worktime each year due to injury on the job. The duration figures can be interpreted as days lost per nonfatal injury that results in any loss of worktime of one day or more. As the data make clear, most of the variation in expected days lost across industries results from differences in the incidence of injury: Incidence and duration have similar variances, even though duration has a much higher mean. Nonetheless, there is substantial variation in the duration of injuries among industries, and thus substantial room for that variation to allow us to test the hypotheses we have discussed.[7] Incidence and duration are also far from

Table 1

Variable Means, Injury Data Linked to PSID Data [a]

Year	**1980**	**1981**
LW	79.82 (57.05)	76.01 (53.79)
INC	4.87 (3.07)	4.64 (2.88)
DUR	15.77 (3.32)	15.64 (3.31)

[a] Standard deviations of the means in parentheses here and in Table 6.

perfectly correlated: Among the observations for 1980 the correlation is only +.30; for 1981 the correlation is only +.32.

The wage measure used in the various specifications is the hourly wage rate on the worker's main job. For salaried workers the PSID bases this measure on the worker's salary divided by some standard working hours. The vector X in (16) and (17) is specified to include a number of measures that have become quite standard in the literature. Thus linear and quadratic terms in total full-time experience since age 18, and in years of tenure with the employer, are included in the equation, as are years of schooling completed. Demographic variables--race, sex, and marital status--are also included, as are indicators of the worker's union membership, region (South) and city size. Also included are weeks worked and hours worked per week in the previous year (1979 or 1980), measures designed to control for the different average rates of pay produced by overtime premia, lower wage rates for part-time workers, etc. Finally, in some of the estimates dummy variables for five major industries--durable manufacturing, nondurable manufacturing, agriculture and mining, transportation, communications and public utilities, and wholesale and retail trade--are also included.[8]

Table 2 presents the parameter estimates for the sample observed in 1980, and Table 3 shows the estimates for 1981. In each table the results are presented without and with the inclusion of the vector of one-digit industry dummy variables. The results of estimating (16) with $\alpha_2=0$, and (17) with $\alpha_2'=1$, are quite consistent with those produced by a number of earlier studies. Since some control for industry differences not attributable to differences in injury rates is probably desirable, most of the remaining discussion refers to parameters estimated in the presence of the industry dummy variables. Even when industry is controlled, though, there is a

Table 2

Parameter Estimates, PSID Data, 1980, Linked to 1980 Injury Data [a]

Equation:	α_1 or α_1'	α_2 or α_2'	$\bar{R}^2$
No Industry Dummies			
(16)	.001199 (7.31)	0	.5181
(16)	.000906 (1.94)	.000014 (.67)	.5179
(17)	.0858 (7.45)	1	.5188
(17)	.3486 (8.21)	.1368 (3.72)	.5315
Industry Dummies			
(16)	.000648 (3.70)	0	.5599
(16)	-.000062 (-.13)	.000033 (1.64)	.5604
(17)	.0479 (3.48)	1	.5594
(17)	.3166 (7.26)	.0564 (1.25)	.5714

[a] t-statistics in parentheses here and in Tables 3, 4, 7, 8 and 9.

Estimating equations also include education, linear and nonlinear terms in experience and tenure, region, marital status, weeks and hours in 1979, sex, race, union status, city size, and three occupation dummies.

Table 3

Parameter Estimates, PSID Data, 1981, Linked to 1981 Injury Data [a]

Equation:	α_1 or α_1'	α_2 or α_2'	$\bar{R}^2$
No Industry Dummies			
(16)	.000821 (4.85)	0	.4980
(16)	.000244 (.50)	.000028 (1.25)	.4981
(17)	.0485 (4.49)	1	.4970
(17)	.2860 (6.50)	.0384 (.87)	.5058
Industry Dummies			
(16)	.000268 (1.45)	0	.5337
(16)	-.000504 (-1.02)	.000037 (1.68)	.5342
(17)	.0135 (1.04)	1	.5333
(17)	.2569 (5.55)	-.0568 (-1.05)	.5414

[a] Estimating equations also include education, linear and nonlinear terms in experience and tenure, region, marital status, weeks and hours in 1980, sex, race, union status, city size, and three occupation dummies.

noticeable positive effect of increased expected lost workdays on wage rates.

The estimates of (16) in which α_2 is free to vary do not give very satistactory results. The increase in the $\bar{R}^2$ is very small, and neither coefficient is significant at conventional levels when the industry dummy variables are included. Apparently the collinearity between LW and LW•DUR is causing problems.[9] When the less restricted equation (17) is estimated we see striking evidence that duration and incidence of injuries do not produce the same compensating wage differentials. The effect of increased duration, α_1', is highly significant and positive in both years; that of increased incidence, $\alpha_1'\alpha_2'$, is positive one year, negative the other, and insignificant in both.[10] Reestimates of (16) and (17) on samples of blue- and white-collar workers separately yield slightly weaker results; but the same qualitative result, a significantly greater impact of duration than of incidence on wages, exists in each of these subsamples too.

One might argue that the compensating differential will be affected by the extent to which workers are insured against the wage loss by social legislation. Thus while many studies ignore this issue, some--Arnould-Nichols (1983), Butler-Worrall (1983) and Ruser (1985)--include replacement rates under workers' compensation benefits in equations like those presented in Tables 2 and 3. Accordingly, the equations were reestimated with various replacement rates included.[11] Adding these measures had no significant effect on the other parameter estimates. Also, the replacement rates usually did have the expected negative coefficients, but these were never significantly negative.[12] Our focus on injuries may explain the departure of these results from those of Arnould-Nichols (1983) on workplace fatalities. Workers' compensation is not an entitlement program; it has long waiting periods, and its receipt is uncertain in the case of most injuries.[13] Thus it perhaps should not be surprising that it does not affect the size of compensating differentials for workplace injuries.

There is some evidence (Duncan-Stafford, 1980) of a link between compensating wage differentials for workplace hazards and the union relative wage effect. That study indicates that part of the union wage advantage represents compensation for exposure to risks in the workplace. To examine whether the obverse is true, as Viscusi (1979) indicates, and, in particular, whether unions have different impacts on the compensating differentials for incidence and duration of risk, we respecify (17). One respecification, which constrains the effects of duration and incidence to be equal, replaces the terms in lnDUR and lnINC in (17) with:

(18) $\alpha_3 \ln LW + \alpha_3' \, UN \cdot \ln LW$

where UN=1 if the worker is a union member.[14] The second respecification allows duration and incidence to have different effects on wages in union and nonunion employment by respecifying (18) as:

(19) $\alpha_4 \ln DUR + \alpha_4' \, UN \cdot \ln DUR + \alpha_5 \ln INC + \alpha_5' \, UN \cdot \ln INC$

The results of estimating equations based on the specifications in (18) and (19) are shown in Table 4 for both 1980 and 1981. The estimates of (18) indicate that, in our linked micro--industry data as in Viscusi's (1979) estimates based on self-reported risks, unionized workers receive an extra compensating differential for risks on the job. This can be interpreted as showing that the informational effects of unions produce increased compensation for what would not be as clearly perceived by workers negotiating individually. The estimates of (19) demonstrate that this compensation is almost entirely for increases in the incidence of the loss: The interaction term between union status and duration is very small, while that between union

Table 4

Parameter Estimates, PSID Data, 1980 and 1981, Linked to Injury Data, With Interaction Terms in Union Membership [a]

	1980		1981	
Equation:	(18)	(19)	(18)	(19)
α_3 or α_4	.0693 (6.14)	.3504 (8.55)	.0325 (3.06)	.2935 (6.88)
α_3' or α_4'	.0440 (9.99)	.0152 (1.06)	.0458 (10.26)	.0139 (.97)
α_5	---	.0147 (1.04)	---	-.0230 (-1.66)
α_5'	---	.0971 (4.01)	---	.1052 (4.28)

[a] The vector of dummy variables for 1-digit industry is included in the equations, as are the variables listed in the notes to Tables 2 and 3.

status and incidence is large and signficant. Our findings suggest that changes in the incidence of losses are not well perceived by workers, while changes in their duration are; if this is so, then the different impacts of unionism on these compensating wage differentials are consistent with the view of unions as organizations that increase workers' awareness of, and rewards for, poorly perceived, generally applicable risks in the workplace.

As another way of examining the differential impacts of incidence and duration of workplace injuries on wages, we calculate the wage-incidence and wage-duration elasticities for both samples. These estimates, based on the unconstrained versions of (16) and (17) that include one-digit industry dummy variables presented in Tables 2 and 3, are listed in Table 5. They show very clearly that the positive effect of injury rates on wages is produced by the duration of the injury. An increase in the risk of injury, holding duration constant, produces only a very slight compensating wage differential. An increase in duration, holding the risk of injury constant, produces a much larger effect on wages. This is especially true if the elasticities are based on (17), which allowed the effects to vary more freely and which produced the higher $\bar{R}^2$ in both years.

Dorsey's (1983) estimates using establishment data show significant impacts of both the incidence and duration of nonfatal injuries. However, using his published means and estimated equation describing lnW, we calculate from his equation a duration elasticity of .28, and an incidence elasticity of .11. The similarity of these elasticities to those based on equation (17) is remarkable given their totally different underlying sources of data and econometric specification.

To examine the importance of the differential effects of incidence and duration on wages, consider what would happen if the average duration of

Table 5

Elasticities of Wage Rates With Respect to

Duration and Incidence of Injuries [a]

	Based on Equation:	**1980**	**1981**
$\partial \ln w / \partial \ln DUR$	(16)	.0868	.0503
	(17)	.3166	.2569
$\partial \ln w / \partial \ln INC$	(16)	.0442	.0049
	(17)	.0179	-.0146

[a] All the elasticities are based on the estimates in Tables 2 and 3 in which the vector of industry dummies is included.

nonfatal injuries dropped by two standard deviations, while the mean expected time lost remained unchanged because of an offsetting increase in incidence. Using estimates based on (17) and on data from 1980, we calculate that the average worker would pay 15 percent of the current average wage to obtain a change in working conditions that would alter outcomes in this manner; using the 1981 estimates, the wage-equivalent of the utility gain implicit in this change is 14 percent. Clearly, there are potentially substantial gains in welfare from reducing the duration of workplace injuries.

IV. An Application to Unemployment

There are two distinct strands in the literature on compensating differentials for the risk of unemployment. One (Hall, 1972; Topel, 1984), examines how wages differ across high- and low-unemployment industries and regions at a point in time, and thus presumably measures the extent of compensating differentials for long-term (structural) differences in unemployment. The other (Abowd-Ashenfelter, 1981) examines how wages differ across industries and occupations with varying probabilities of cyclical unemployment. Clearly, the two strands are distinct in terms of empirical specification (though Li, 1986, provides an initial attempt to estimate both in the same model). In terms of their relation to the underlying theory, though, no such distinction exists. Both types of differential presumably arise out of workers' awareness that there are occupational and industrial differences in the risks of both types of unemployment. That being the case, our theory suggests that we should observe compensating differentials for both risks, and that in each case the differential should be greater, given identical expected losses, for increases in the duration of loss than for increases in its incidence.

To examine the hypothesis in this context, we again use data from the

Panel Study of Income Dynamics, in this case only from the 1981 interviewing wave. The wage and background data are as in Section III; however, because we did not require that wage data be available for two consecutive years, and because the link to unemployment data was possible for all industries, 2625 observations are available for this part of the study.

The unemployment data are based on supplementary questions on work experience appended to the March Current Population Survey. Because we wish to examine compensating differences for both structural and cyclical unemployment, we use data for both 1979 (a cyclical peak) and 1982 (a cyclical trough), data from the March 1980 and March 1983 CPS.[15] For workers interviewed in March of the subsequent year, data are provided on the fraction experiencing some unemployment and on the distribution of weeks of unemployment among those individuals. Workers' affiliations by two- or three-digit industry are based on where they worked the longest during the calendar year (not where they worked at the date of the interview).

Because the duration data are categorical, it was necessary to aggregate them using some assumptions about their distributions within the categories.[16] We assumed that the hazard rate of leaving unemployment was constant within each category, and that the fraction of workers remaining unemployed at the end of an interval equalled the published fraction remaining unemployed. This technique produced the data on duration and incidence for 1979 and 1982 and for the peak-to-trough variation, all of which are presented in Table 6.

There is much greater variation across industries in the incidence of unemployment than in its duration. This is true in both a peak year, 1979, and at a business-cycle trough, 1982. Moreover, even though the decomposition of the cyclical increase in unemployment into cyclical changes in duration and incidence shows that both increased roughly equally, the variance in the

Table 6

Variable Means, Unemployment Data Linked to PSID Data, 1981

	Year of Unemployment Data		
	1979	1982	Change (in logs)
Unemployment Rate (percent)	3.60 (1.71)	6.48 (3.14)	.57 (.24)
INC (percent of workers)	15.91 (6.58)	22.22 (9.79)	.31 (.21)
DUR (weeks)	11.59 (1.27)	14.99 (1.31)	.26 (.09)

cyclical change in incidence across industries was much greater than that in duration. The greater variation in incidence than in duration was the same phenomenon that we observed in workplace injury rates. Also as in those data, the simple correlations between duration and incidence are not particularly high: For 1979 and 1982 they are +.32 and +.33 respectively; for the cyclical changes in duration and incidence, the simple correlation is only +.15.

A. Unemployment in 1979

The results of estimating variants of equations (16) and (17) over the 1981 PSID data linked to the 1979 work-experience data are presented in the first four rows of Table 7. In this table and in Table 8 only estimates based on equations that include the vector of one-digit industry dummy variables are presented. (The results do not differ qualitatively when this vector is excluded.) The results are very disappointing. There is a negative and significant relation between the wage rate and the unemployment rate of experienced workers in the industry (as shown in the constrained versions of equations (16) and (17)). Indeed, as the unconstrained version of (17) shows, it is differences in the duration of unemployment in 1979 that are most strongly linked to (lower) wage rates; the effect of incidence is smaller and not significant.

Two explanations for these unexpected results were explored. We saw in Section III that interstate differences in worker's compensation benefits did not affect compensating differentials paid to workers in different industries. However, unemployment insurance is a more widespread transfer than is worker's compensation; and, more important, it is a transfer that will be received with near-certainty should a particular loss occur. Thus, if interindustry differences in duration and incidence are correlated with the

Table 7

Parameter Estimates, PSID Data, 1981, Linked to 1979 Unemployment Data

Including One-Digit Industry Dummy Variables [a]

Equation:	α_1 or α_1'	α_2 or α_2'	Potential Duration	Percent Long-term Unemployed	$\bar{R}^2$
(16)	-.0174 (-2.55)	0			.5247
(16)	.1164 (4.58)	-.0101 (-5.47)			.5299
(17)	-.0622 (-2.89)	1			.5251
(17)	-.4631 (-5.99)	.0024 (.05)			.5301
(17)	-.4606 (-5.96)	.00553 (.11)	-.00598 (-1.93)		.5306
(17)	.5217 (3.33)	-.1132 (-2.34)	-.00613 (-2.00)	-.0316 (-7.18)	.5396

[a] Also included are the same variables that were included in the regressions presented in Tables 3 and 4, and in Tables 8 and 9.

generosity of state UI programs, failure to include some measure of the latter will bias estimates of compensating differentials for the risk of unemployment. To examine this possibility we linked the state average potential duration of regular UI benefits to the data on household heads from the 1981 PSID.[17]

The result of adding the potential duration of regular UI benefits in the state in which the worker resides to (17) are shown in the fifth row of Table 7. Workers in states that offer UI benefits with longer potential duration do receive lower wages, with each extra week of potential duration reducing wages by .6 percent. However, inclusion of the UI measure does not qualitatively affect the estimated impacts of duration and incidence. By inference there is little correlation across workers between interstate differences in the generosity of UI and interindustry differences in unemployment duration and incidence.

The second explanation is that interindustry differences in the mean loss are unimportant, and that workers require compensation only for the risk of a long-duration loss (since only that loss will not be at least partly compensated by UI benefits). This view is consistent both with the derivation in Section II and with the notion that the value of leisure during the first part of a spell of unemployment is quite high. To examine this possibility we also added the percentage of experienced workers by industry who were unemployed more than 26 weeks to the estimating equations.

This addition produced some interesting changes in the results. Looking at the sixth row of Table 7, one sees that increases in the average duration of unemployment produce the expected positive effect on wages, while greater incidence still reduces wages (though only slightly). However, the largest effect is the very significant negative impact of increases in the percentage of long-term unemployed. The results thus still confound the predictions of our simple model.

The estimates indicate clearly that longer-duration unemployment, especially increases in the fraction of the unemployed in an industry in the upper tail of the distribution of spells by length, is associated with lower wage rates. One explanation for this apparent anomaly is that some individuals move frequently between employment and nonparticipation, much of the latter of which is recorded as unemployment, because they have reservation wages that are high relative to their market wages. Unless differences in market wages are associated in the population with even larger differences in reservation wages, these individuals will tend to be those with below-average market wages. Thus an industry recording a large amount of long-duration unemployment may also employ workers who command lower-than-average wages. The spells of long-duration unemployment in such industries will not be compensated by wages because they represent leisure that is valued. This view is supported by inspection in the data we use: The industries having the largest percentages of long-term unemployed among experienced workers attached to the industry in 1979 were private household services, welfare and religious services, and agriculture. The lowest percentages were in automobile manufacturing, apparel manufacturing, and stone, clay and glass manufacturing.

B. Cyclical Changes in Unemployment, 1979-82

Equations (16) and (17) were reestimated using the cyclical changes in (the logarithms) of the duration and incidence of unemployment by industry. The results of this estimation are shown in Table 8. As the first four rows show, compensating differentials exist for cyclical variations in the

Table 8

Parameter Estimates, PSID Data, 1981, Linked to 1979-82 Unemployment Changes, Including One-Digit Industry Dummy Variables

Equation:	α_1 or α_1'	α_2 or α_2'	Potential Duration	Cyclical Change in Percent Long-term Unemployed	$\bar{R}^2$
(16)	-.0091 (-1.68)	0			.5241
(16)	-.0400 (-3.79)	.00694 (3.42)			.5260
(17)	.0307 (.91)	1			.5237
(17)	.4392 (5.28)	-.1060 (-1.28)			.5287
(17)	.4376 (5.27)	-.1024 (-1.22)	-.0059 (-1.91)		.5292
(17)	.0098 (.07)	-7.215 (-1.91)	-.0056 (-1.80)	.0148 (3.49)	.5312

incidence and duration of unemployment that are remarkably like those that we demonstrated in Section III exist for workplace injuries. The estimates of the unconstrained version of (17) indicate that the compensating differential is paid only for differences in cyclical changes in duration; cyclical changes in incidence have no impact on wage differentials once changes in duration are accounted for.[18] This conclusion is underscored by a comparison of the $\bar{R}^2$ in the regressions in the third and four rows.[19]

As in Section IV.A., we added a measure of the generosity of unemployment benefits, the potential duration of benefits, to the equations. Also, the cyclical change in the percentage of long-term unemployed workers by industry was added. Examining the fifth row in Table 8, we again find that workers living in states with a longer average potential duration of UI benefits receive lower wage rates. The results of including the cyclical change in the percentage of long-term unemployed are striking. One notes from the last row in the table that there is no independent impact either of the change in the average duration or of the change in incidence. Rather, there is a very significant positive effect of the change in long-term unemployment. The positive compensating differential that exists for larger cyclical changes in the duration of unemployment is entirely due to differentials that are paid in industries where the risk of long-duration unemployment increases most during recessions. Since it is precisely the cyclical increase in long-duration unemployment that is not automatically compensated by unemployment insurance, this result makes sense.

Since we know unions affect cyclical changes in employment (see Medoff, 1979), it is worth examining how they affect the compensating differentials that we have demonstrated exist for cyclical changes in unemployment duration. The results of estimating (18) and (19), the versions of (17) respecified to include interactions of union status with expected loss, and with duration and incidence separately, are shown in Table 9.[20] The most striking result is that the interaction term involving lnDUR is highly significant and negative. Indeed, the estimates of (19) demonstrate that the large positive effect of greater cyclical changes in unemployment duration on wages arises solely in nonunion employment; among unionized workers the effects are negative and insignificant. (A test of the joint significance of α_4' and α_5' yielded $F(2, 2598) = 4.69$, significant at the 99 percent level of confidence.) The results of estimating (19) show that a failure to understand that workers react more strongly to differences in duration than in incidence would have prevented one from seeing how unions affect these compensating differentials: If changes in duration and incidence are constrained to have the same effect, the interaction term with union status is not significant.

The lack of a compensating wage differential in unionized employment for differences in the cyclicality of the duration of unemployment is consistent with several models of union behavior. One standard analysis assumes that unions seek to maximize the utility of the median member (voter). Assume also that demand is not so highly variable over the cycle that the worker with the median amount of seniority will be laid off during a recession. That being the case, the existence of larger variations in the cyclicality of unemployment will not affect union bargainers' wage policy, as the median union member will be unconcerned about such variations.

Table 10 presents estimates of the elasticity of wage rates with respect to interindustry differences in cyclical changes in the average duration and incidence of unemployment. The elasticities are based on the estimates of the unconstrained versions of equations (16) and (17) (excluding the measures of potential duration of UI benefits and of long-term unemployment). The

Table 9

Parameter Estimates, PSID Data, 1981, Linked to 1979-82 Unemployment Changes, with Interaction Terms in Union Membership

Equation:	(18)	(19)
α_3 or α_4	-.0164 (-.52)	.2955 (3.88)
α_3' or α_4'	-.0187 (-.38)	-.3530 (-2.54)
α_5	---	-.1453 (-3.41)
α_5'	---	.1075 (1.66)

Table 10

Elasticity of Wage Rates With Respect to Cyclical Changes in the Duration and Incidence of Unemployment, 1979-1982

	Based on Equation: (16)	(17)
$\partial \ln w / \partial \ln DUR$	.1523	.4392
$\partial \ln w / \partial \ln INC$	-.3529	-.0467

elasticities based on (17) are quite similar in magnitude to the elasticities presented in Table 5; they suggest a huge compensating differential for changes in duration, with essentially no compensation for cyclical changes in the incidence of unemployment. Using these estimates, we calculate that a two-standard-deviation decrease in the cyclicality of duration that is accompanied by an offsetting increase in the cyclical variability of incidence would induce an 8 percent decrease in wage rates. As with workplace injuries, there is evidence that there would be substantial welfare gains to reducing the cyclical variability of unemployment duration.

V. Conclusions

We have derived a model of compensating differentials for wage losses which recognizes the importance of risk aversion. The model predicts that wage differentials will respond more strongly to an increase in the duration of the wage loss than to a rise in its incidence that produces an equal increase in the expected loss. This prediction was first verified using two cross sections of data on individuals' wages and characteristics linked to aggregate data on the injury experience of the three-digit industries in which they work. We found that most of the compensating differential for higher nonfatal workplace injuries stems from the large negative effect of an increase in the expected duration of an injury on the wage. The hypothesis was then examined in the context of compensating differentials for unemployment, both structural and cyclical. There was no support for it in cross-section data, perhaps because of unmeasured differences across industries in reservation wages. However, we found that the compensating wage differential for differences in the cyclicality of unemployment is mainly a result of compensation for differences in the cyclicality of unemployment duration. Moreover, the elasticities of wage rates with respect to these differences are quite close to those with respect to differences in the duration of injuries.

We have also shown that union intervention in the process that generates compensating differentials for wage losses differs by the type of loss. In the case of losses due to injuries unions raise the compensation for increases in the expected incidence of the loss and have little impact on compensation for increased duration. This contrasts to our finding that unionized workers receive little compensation for the risk of cyclical unemployment, especially cyclical increases in unemployment duration. Upon first glance these results appear quite contradictory. If one considers the nature of the losses involved, though, the two sets of results are completely consistent. Unlike the risk of cyclical unemployment, which is borne in most cases by junior workers, the risk of injury affects all workers in a plant; there is very little unions can do to shift the risk away from the median member. Thus we should expect that unions will bargain for higher wages to compensate the median member for the risk of injury, while they are less concerned with the effects of a higher risk of cyclical unemployment, especially longer-duration cyclical unemployment, that do not affect most members.

In a world of complete information and certain receipt of employer-financed insurance for wage losses there is an equivalence between the cost of compensating wage differentials to induce workers to accept risks and the cost of social insurance. That equivalence breaks down if, as our results indicate, workers' risk aversion leads them to demand extra compensation for increases in the duration of loss beyond that which compensates them for the expected loss. The cost of insurance would merely equal the expected loss, while compensating differentials will vary depending on the relative sizes of the two components of that expected loss. In general, then, the marginal benefit to the employer of greater safety will depend upon the absence or presence of legal insurance requirements. Our results therefore indicate that, because of the special roles of duration and risk aversion, welfare depends upon the insurance regime assumed.

While the empirical research on workplace injuries and cyclical unemployment offers evidence supporting our predictions about the importance of duration of loss in producing compensating wage differentials, it is only the beginning of research on this issue. Additional work on alternative sets of data is needed. Also, further work should test the various explanations for the existence of especially large compensating differentials for the duration of loss.

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FOOTNOTES

1. The only theoretical discussion of this issue is in Adams (1985), who only notes that compensating differentials for unemployment could differ depending on whether the increased likelihood of loss is due to greater incidence or longer duration.

2. Multiple tangencies between the wage frontier and worker indifference curves are possible, even if indifference curves are all identical, because the wage frontier is responsive to market demands for products requiring varying risks of wage loss.

3. If the worker perceives that wages are reduced by a fraction ϵ of the expected loss $\beta\gamma$ in order to help to finance wage replacement, and if w_0 represents the worker's true marginal product, then the second term in (15) becomes $\beta\gamma[\quad + \epsilon]$.

4. One would also like to test the equations using job changers, those for whom fixed effects can be removed. This would, as Duncan-Holmlund (1983) point out, reduce biases in the estimates of compensating differentials, though the reduction is less in our data than it would be if we used self-reported risks. Unfortunately, 1981 was the first year the necessary detail on industry affiliation was given in the PSID.

5. The data are from BLS, Occupational Injuries and Illnesses in the United States by Industry, 1980, 1981, Bulletins 2130 and 2164.

6. Clearly, since the data cover industries and the observations are on individuals, there is no simultaneity problem. There is, however, a potential problem of truncation of the duration data, as only days lost during the particular year are included in the calculation of DUR for an industry. Unfortunately, without additional information it is impossible to tell whether this measurement problem produces different biases on the estimates of the separate effects of duration and incidence.

7. The ranges of LW in 1980 and 1981 are from 3.2 to 338.9, and 2.3 to 289.3; the ranges of INC are from .2 to 14.9, and from .2 to 14.4; those of DUR are from 9.0 to 35.75, and from 9.0 to 37.57.

8. One could specify a finer breakdown by industry. However, a complete set of dummies, one for each three-digit industry, would wipe out the coefficients on the injury variables, since these are available only at that level of disaggregation. There is thus an inherent problem in this and all other studies of compensating differentials for risk of workplace injury or fatality that use micro data linked to industry or occupation statistics: One cannot completely distinguish the effects of other industrial or occupational characteristics that are correlated with the incidence and duration of injury and that affect wage differentials from those of the injury hazards themselves.

9. That this is producing the difficulty is suggested by the relative lack of

variation in DUR that we noted above.

10. While we have used logarithmic forms of DUR and INC here, qualitatively similar results are produced when linear forms are included in a respecified version of (17).

11. Richard Butler kindly provided the data used in his study. For observations in the 35 states with adequate data, equations including replacement rates under workers' compensation were estimated.

12. Our finding of little effect of workers' compensation benefits on the compensating differentials parallels that of Ruser (1985). The difference between our results and those of Butler-Worrall (1983) may stem from their use of aggregate wage data.

13. Although disturbingly little information is available on actual benefits paid, as opposed to benefit schedules, the evidence does suggest the haphazard nature of income replacement. Thus Interdepartmental Workers' Compensation Task Force, <u>Research Report</u>, Volume VI, 1981, shows that actual replacement rates for varying degrees of permanent partial disability ranged from .45 to 1.53 in Wisconsin, and from 1.85 to 13.85 in Florida.

14. Union membership rather than collective bargaining coverage is also used in studies of compensating differentials by Viscusi (1979), Olson (1981) and Duncan-Stafford (1980).

15. The March 1980 data are unpublished and were kindly provided to us by Paul Flaim; the March 1983 data are presented in Table B-12, BLS, <u>Work Experience of the Population in 1981-82</u>, Bulletin 2199, 1984.

16. The data are divided into durations of 1-4 weeks, 5-10, 11-14, 15-26, and 27 plus. The duration data measure the total weeks of unemployment experienced during the previous year. Thus an individual with two ten-week spells would be recorded as having unemployment with a duration of twenty weeks. Also, as with the injury data used in Section III, reported spell duration may be truncated because spells that overlap calendar years are not fully reported. Without knowing more detail about their distribution, though, we cannot tell what are the relative biases to the separate estimates of incidence and duration effects on compensating differentials.

17. The UI data are for 1980, the most recent available, and are taken from Employment and Training Administration, <u>Handbook of UI Financial Data</u>, ET Handbook 394. Topel (1984) found that adding a measure of the replacement rate (the weekly benefit relative to the individual's wage) to equations that showed no compensating wage differential for unemployment changed those results drastically and made the differential significant and positive. However, the replacement measure included the dependent variable in its denominator. Moreover, since our theory is based on aversion to the risk of a long-duration loss, a measure of interstate differences in the extent to which long-duration losses are covered is more appropriate for our purposes.

18. The double recession from 1980-82 was reputed to be especially heavily concentrated in high-wage industries. If this is true, and if the variables in the vector X and the vector of one-digit industry dummy variables do not account for all other factors, it may be that the peculiarities of that

recession are producing our results. Thus the high-wage industries would be those in which duration rose the most, not because high wages represent compensating differentials, but because high wages were associated with, and may even have induced, above-average cyclical increases in unemployment. Without data from additional recessions this possibility cannot be distinguished from our explanation.

19. The result does not depend on our use of logarithmic forms of the measures of duration and incidence: When linear forms were added the results changed little (though the R^2 were slightly lower). Similarly, the results differ little when the sample is restricted to blue-collar workers.

20. The equations presented in the table do not include the measure of cyclical changes in long-term unemployment.