Worker Learning and Compensating Differentials

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The authors hypothesize that in industries with relatively high levels of job-related injury risk, workers with longer job tenure will more clearly appreciate the degree of job risk than will newly hired workers, and will thus be more willing to accept lower wages in return for higher workers' compensation benefits. This hypothesis is confirmed by an analysis of quit behavior using 1981–83 data from the Michigan Panel Study of Income Dynamics and 1981–85 data from the National Institute of Occupational Safety and Health.

In the standard compensating wage differential model, workers value their wage and workers' compensation components based on full job risk information. Market forces generate positive wage differentials as ex ante compensation for exposure to relatively high risk. Similarly, market forces generate wage offsets for the increases in ex post risk compensation embodied in workers' compensation benefits.

These predictions can be modified to take into account potential imperfections in worker information, as in Viscusi (1979a,b, 1980a,b,d), where the role of learning is incorporated into the worker's decision model. The potential for learning about risks introduces a new market response through worker quitting after the acquisition of adverse risk information. In a full information world, after controlling for health status, no unexpected job risk–quit relationship will be observed. In the more realistic sequential decision model in which there is an opportunity for learning, the acquisition of adverse new information by the worker on the job may lead the worker to quit.

With the exception of the experimental results reported in Viscusi and O'Connor (1984), in which worker responses to alternative chemical labels were monitored, tests of the standard compensating differential model and of the learning models have been distinct, as each focuses on a different aspect of labor market behavior. The empirical evidence supporting compensating risk differentials is substantial: greater job risks boost worker wages, and workers are willing to accept a wage cut in return for higher workers' compensation benefits.[1] These results are

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The data and programs used in this study are available on request to Michael J. Moore, Fuqua School of Business, Duke University, Durham, NC 27706.

[1] See, for example, Smith (1979) and Viscusi (1979a) for analysis of wage-risk tradeoffs. Estimates of wage–workers' compensation tradeoffs appear in Arnould and Nichols (1983), Butler (1983), Dorsey and Walzer (1983), and Viscusi and Moore (1987). In

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the main predictions of the standard compensating differential theory, and they continue to hold if learning is introduced. Market tests of the role of worker learning, on the other hand, have focused on two other empirical issues—the effect of injury experiences on workers' risk perceptions and the positive effect of job risks on worker quitting.[2]

The focus of this paper is broader than that of separate analyses of the wage and quit effects of job risks because we use the relationships typically estimated and tested in the standard compensating differential theory to examine the job risk–learning model as well. In particular, using a large data set on workers in the early 1980s, we evaluate the tradeoffs between wages and workers' compensation benefits and between wages and risks implied by worker quit behavior, and compare these tradeoffs across worker tenure groups.

Younger workers will assess job risk less precisely than more senior workers, since their informational base for making these judgments is smaller. In multi-period models that incorporate learning and experimentation with risky jobs, workers' reservation wage rates will be less for jobs posing less precisely understood risks, for any given mean level of risk. The empirical prediction is that more senior workers will demand greater compensation for risk because of their more precise judgments.

In addition, workers who are on the quit margin (those who would need only a small inducement to quit their jobs) will have greater subjective risk perceptions than other workers. These workers consequently will demand greater wage compensation for higher risk levels, since they will be comprised disproportionately of workers whose risk beliefs have been adversely affected by on-the-job experiences. These workers will also assess a greater chance of receiving workers' compensation benefits, so their expected value of higher benefits will be greater. We therefore expect to observe greater wage-risk and wage–workers' compensation tradeoffs for these workers.

Moore and Viscusi (1990) we provide a literature review of estimates of the effects of fatality risks and workers' compensation on wages.

[2] See Viscusi (1979a,b, 1980a,b,d, and 1983) and Viscusi and O'Connor (1984).

Theoretical Framework

The learning model that we apply here to the wage–workers' compensation tradeoff was first introduced in Viscusi (1979a,b, 1980a,b,d).[3] Our overall objective is to explore the relationships between the wage-risk tradeoff and the wage–workers' compensation tradeoff for new workers and for senior workers on the quit–no quit margin.

The essential ideas can be captured in a two-period model. Let there be two health states: healthy and injured. In the good health state the worker receives a wage rate w, from which he derives utility $U^1(w)$. In the injured state the worker receives workers' compensation benefits b, where $w \geq b$. This assumption reflects the structure of workers' compensation programs in virtually every state, since benefits are typically two-thirds of the wage or less, except for workers with very low wages. We assume that the worker would rather be healthy than not (that is, $U^1(x) > U^2(x)$), has a higher marginal utility of income when healthy ($U_x^1 > U_x^2$), and is either risk-averse or risk-neutral ($U_{xx}^1, U_{xx}^2 \leq 0$). The worker values payoffs over time using a discount factor β that equals the reciprocal of 1 plus the interest rate.

Suppose that there are two possible jobs, a risky job and a safe job. We can assume with no loss of generality that the safe job poses no risk of injury.[4] The safe job offers a payoff w_0 forever. The risky job offers the worker an initial perceived probability of not being injured equal to p and a $1-p$ chance of suffering an injury that lasts a single period. If the worker is not injured in period 1, he revises his

[3] A variant of this analysis without learning appears in Diamond (1977).

[4] One could adopt the assumption that the alternative job poses a known risk of injury without altering the model structure, even in the n-period case. If both jobs are uncertain and there are more than two periods, the model structure becomes more complex. See Viscusi (1979a) for these extensions.

assessed probability of not being injured upward to p^+. If he is injured, the assessed probability is p^-. It is also possible, as noted by Viscusi (1979a), that workers revise their expectations based on observations of other workers' injury experiences. The revision of workers' probabilistic beliefs follows a standard Bayesian learning process, where

$$p^+ > p > p^-. \tag{1}$$

The workers' initial job decision involves a choice between two periods of work on the safe job or initial work on the risky job, after which he can quit if he is injured in period 1. As first noted by Viscusi (1979a), the worker's problem mirrors the classic two-armed bandit problem, which describes the optimal sequence of plays on two slot machines. On one machine the probability of success is known, and on the other it is uncertain. Thus, a payoff on the uncertain machine yields information in addition to monetary rewards. For this class of two-armed bandit problems, it is shown in Viscusi (1979a,b, 1980a,b,d) that the stay-on-a-winner rule is always optimal. The worker will not leave the risky job after a favorable experience in period 1. The worker also will not leave the safe job once he starts on it.

The wage package for the marginal new worker attracted to the firm must satisfy the condition that expected lifetime utility, V, is equal between the two jobs, given the opportunity the worker has to switch from the risky job following an unfavorable period 1 outcome:

$$\begin{aligned} V &= U^1(w_0)\,(1+\beta) \\ &= pU^1(w) + (1-p)U^2(b) \\ &\quad + \beta p[p^+U^1(w) + \\ &\quad (1-p^+)U^2(b)] \\ &\quad + \beta(1-p)\text{Max}[U^1(w_0), \\ &\quad p^-U^1(w) + (1-p^-)U^2(b)]. \end{aligned} \tag{2}$$

If we set $U^1(w_0)$ equal to zero, with no loss of generality, we have

$$\begin{aligned} V = 0 &= pU^1(w) + (1-p)U^2(b) \\ &\quad + \beta p[p^+U^1(w) + \\ &\quad (1-p^+)U^2(b)] \\ &\quad + \beta(1-p)\text{Max}[0, p^-U^1(w) \\ &\quad + (1-p^-)U^2(b)]. \end{aligned} \tag{3}$$

Not all workers will quit their jobs prior to period 2 after an unfavorable period 1 job experience. The focus here, however, is on the wage-benefit tradeoff of the marginal senior worker relative to the pay package that will attract the worker to the job initially. The marginal senior worker will quit after an adverse experience in period 1, so the wage package (w,b) sufficient to attract the worker initially must satisfy

$$\begin{aligned} V = 0 &= pU^1(w) + (1-p)U^2(b) \\ &\quad + \beta p[p^+U^1(w) + \\ &\quad (1-p^+)U^2(b)], \end{aligned} \tag{4}$$

since the last term in equation 3 equals zero after an unfavorable period 1 experience.

The first issue analyzed is the wage–workers' compensation tradeoff that will be reflected in the (w,b) package for new hires. Implicit differentiation of equation 3 yields

$$\frac{\partial w}{\partial b} = \frac{-V_b}{V_w} = \frac{-U_x^2\,[(1-p) + \beta p(1-p^+)]}{U_x^1[p + \beta p p^+]}. \tag{5}$$

The value of $\partial w/\partial b$ represents the wage offset in response to b for the new hire in a two-period job choice problem. The initial wage package (w,b) will be adequate to retain the worker if his on-the-job experiences are favorable.

The worker on the margin at the start of period 2 has an expected utility Z equal to

$$Z = 0 = p^-U^1(w) + (1-p^-)U^2(b), \tag{6}$$

since he is indifferent between leaving (where $U^1(w_0) = 0$) and staying on the risky job. The wage–workers' compensation tradeoff for this worker equals

$$\frac{\partial w}{\partial b} = \frac{-Z_b}{Z_w} = \frac{-(1-p^-)U_x^2}{p^-U_x^1}. \tag{7}$$

One issue that we investigate empirically is the relative magnitude of the wage–workers' compensation tradeoffs in equations 5 and 7. Workers who have experienced or observed an on-the-job injury should value workers' compensation more highly, since they will assess a higher probability of receiving such benefits than will other workers. For workers with an adverse job experience, the expected amount of workers' compensation benefits will be $(1-p^{+})b$. For new hires who plan to quit if their initial period job experience is unfavorable, the discounted expected benefit amount is $(1-p)b + \beta(1-p)(1-p^{-})b$, where these benefits are provided over $(1-p) + \beta(1-p)$ periods. The lower expected amount of benefits per period of work for new hires is reflected in a lower expected utility as well, which will influence the compensating differential they are willing to accept for these benefits. In particular, one would expect

$$\text{(8)}\quad \frac{\partial w}{\partial b}:\ \underset{\textit{Quit Margin Tradeoff}}{\frac{-Z_b}{Z_w}} < \underset{\textit{New Hire Tradeoff}}{\frac{-V_b}{V_w}},$$

or

(9)

$$-\frac{(1-p^{-})U_x^2}{p^{-}U_x^1} < \frac{-U_x^2[(1-p) + \beta p(1-p^{+})]}{U_x^1[p + \beta p p^{+}]}$$

After some algebraic manipulation, equation 9 reduces to

$$\text{(10)}\qquad p^{-} < p^{+}.$$

Given the restrictions on probabilities outlined in equation 1 above, equation 10 always holds.

The second empirical concern will be the effect of worker learning on the character of the wage premiums commanded by job risks. From the development above, equation 4 implies that, for new hires,

(11)

$$U^1(w) = \frac{-(1-p)U^2(b) - \beta p(1-p^{+})U^2(b)}{p\,(1+\beta p^{+})}.$$

Similarly, solving for $U^1(w)$ from equation 6 for workers on the quit margin yields the condition that

$$\text{(12)}\qquad U^1(w) = \frac{-(1-p^{-})U^2(b)}{p^{-}}.$$

The utility metric defined by setting $U^1(w_0)$ equal to zero leads to a negative value of $U^2(b)$, implying that the right sides of equations 11 and 12 are positive. The condition that must be met for the risk premium required by senior workers on the margin to be higher than that of new hires is that $U^1(w)$ be greater for this group. Thus, the right side of equation 12 must exceed the expression on the right side of equation 11. This requirement is always satisfied if inequality 10 holds, as is assumed.

In addition to compensation for risk, the overall wage structure of the firm will, of course, also include returns to worker experience and performance. This wage structure can be viewed as defining the pecuniary returns to the worker over time. As is standard in agency theory models of wages, it is the new hires and other workers on the margin of leaving the firm who are of greatest concern. By altering the entering wage level, the firm ensures a flow of new workers to the firm. Higher wages also diminish the tendency to quit, but since quitters tend to be workers with particularly adverse job experiences, learning-induced quits will continue to occur. Because of the difference in risk perceptions of the potential quitters, both safety and workers' compensation should be more highly valued by this group than by the new worker group, which has a lower assessment of the job risk both for the initial period of work and for their expected duration of work at the firm.

The principal predictions that we explore below stem from the effect of worker experiences on risk perceptions. For senior workers on the margin, the wage–workers' compensation tradeoff should be more negative (and less than zero), as indicated by equation 8. This relationship arises because the higher probability these workers attach to receiving benefits as a result of their on-the-job learning about the risks should be reflected in the preferences captured in quit equations. In addition, the compensating differential experienced workers demand for risk will also rise for any given level of perceived risk, to the extent that more experienced workers place less value on future benefits of job experimentation because of their more precise evaluations of job risk. Experienced workers on the quit margin will also consist disproportionately of workers who have had or observed adverse job experiences, boosting the required wage-risk tradeoff.

The linkage of the theoretical predictions to the empirical model will be fairly direct. Wage and workers' compensation levels are directly observable. We do not have direct measures of workers' subjective probability of being injured for the data set that we will use below, but we do have relatively good data on the fatality rate for the worker's industry. The underlying assumption is that workers will have higher risk perceptions and will be more likely to acquire adverse job information that will generate quit behavior in industries with high objective measures of risk. The results in Viscusi (1979a) and Viscusi and O'Connor (1984) indicate that the job risk–quit relationship is similar whether one uses objective industry risk variables or subjective risk assessments. We use the objective measure here because our focus on state differences in workers' compensation requires that we use a large national data set for which there are no subjective risk data.

Sample and Empirical Results

Data

The main requirement with respect to the survey data for the study is that they include information on wage rates, quit behavior, the worker's state of residence (to establish matchups to workers' compensation benefit formulas), and the worker's state of residence and industry (to establish matchups with risk data).[5] The principal data source we use is the 1981, 1982, and 1983 waves of the University of Michigan Panel Study of Income Dynamics (PSID). The PSID is a broad longitudinal survey data set that contains information pertaining to the characteristics of individuals and their jobs. These data are matched to information on the risk of an on-the-job fatality provided by the National Institute for Occupational Safety and Health (NIOSH) as part of its National Traumatic Occupational Fatality Project (NTOF). Unlike comparable risk data available from other sources, the NIOSH data represent a census of all occupational fatalities, averaged over the years 1981–85. As such, these data are not subject to the sampling error that is present in risk survey data. Furthermore, the NIOSH data vary by both state and industry, making them more comparable to the workers' compensation benefit data, which vary primarily by state. This feature provides a better matchup than the matchup possible using other available national survey data, which vary only by industry. The NIOSH data yield over 400 distinct observations of job risk, thus providing one of the most detailed breakdowns of injury risk currently available.[6]

The workers' compensation data, which are described in detail in Viscusi and Moore (1987), are based on benefits for temporary total disabilities, the injury category under which approximately 65% of total claims fall. Furthermore, in recent years benefit ceilings for temporary total

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[5] An exact matchup of benefits and risk levels with the worker's state of employment, although desirable, is not possible with our data, since sample members report only their state of residence. For the majority of cases, however, the state of residence and the state of employment are the same.

[6] The properties of the NIOSH data are explored in relation to the BLS data in Moore and Viscusi (1988, 1990).

Since the model is based on perceived changes in actual risk, the state-industry risk data will not allow us to distinguish among changes in perceptions, actual risk changes, and changes in risk preferences. We assume that the latter two change slowly, if at all.

disabilities and for permanent total disabilities have become equal in practically all states.[7] Fatality benefits also usually equal benefits for the above two categories, although there are some exceptions. This standardization of benefit ceilings allows a more representative measure of ex post accident compensation than was available in most earlier studies. On the other hand, it ignores some other important aspects of benefit structure, such as differences between ceilings for partial and total disability, waiting periods, and duration. These differences, although important, cannot be captured in a refined fashion, and researchers typically resort to using temporary total disability ceilings or payments as proxies for each state's benefit level.

The workers' compensation benefit levels are matched to workers in the PSID by state. They are then used in conjunction with information on the worker's weekly wage, marital status, and family size to determine the weekly benefit for which the worker qualifies. The benefit variable is computed using the formula b = (2/3 weekly wage) × (1 − D) + (benefit maximum) × D, where D = 1 if the worker qualifies for the maximum benefit level, and 0 otherwise. The variable b is then divided by the worker's after-tax weekly wage to construct the wage replacement rate measure that is used in the empirical analysis.[8]

For purposes of estimation, workers who report their occupation as farming are excluded from the sample, since the agricultural sector is excluded from the NIOSH data. Also excluded are workers whose reported hourly wage is below the statutory minimum wage, non-heads of households, blacks, workers who are over 65 years old or are not in the labor force, and cases with missing data. The sample that remains consists of 2,571 observations on 857 workers. The mix of the workers in the sample follows the expected patterns. Fifteen percent of the sample members are women (FEMALE), 11% are single (SINGLE), and the mean number of dependent children is close to 1 (DEPENDENTS).

The human capital variables include the standard measures. The workers have an average of about 13 years of schooling (EDUCATION), 10 years of experience at their current firm (TENURE), and 18 years of job experience overall (EXPERIENCE). Because of the interrelationships among the various human capital variables, a worker age measure is not included in the analysis. The sample consists primarily of workers in industries in which job hazards are likely to be of consequence. In particular, the sample restrictions we have imposed yield a sample that is over 30% unionized (UNION).

The average after-tax real hourly wage equals approximately $7 in 1982 dollars. The tax component of wages was calculated using information on marginal tax rates provided by the PSID sample members. Although most wage equation studies utilize the pre-tax wage for simplicity, the inclusion of workers' compensation benefits in the equation increases the importance of using after-tax wages, since these benefits have favorable tax status. Moreover, the extent of one's tax savings varies with one's tax bracket and state tax level. Failure to make an adjustment for taxes would thus distort the tradeoff rate between wages and workers' compensation, which is a central empirical concern in this paper.

The death risk measure (RISK), which was described earlier, implies an average death risk of 7.6/100,000 for sample members. This risk level is only slightly different from the national average NTOF risk measure of 9/100,000, so the sample is representative of the industry mix captured in the NTOF data set. The roughly 1/10,000 death risk level should be viewed as a typical risk

[7] A detailed exploration of the differences in disability benefits and their interrelationship is provided by Burton and Krueger (1986) and Krueger and Burton (1983). A detailed analysis of the important permanent partial disability component is provided by Burton (1983). More generally, see Berkowitz and Burton (1987) for an analysis of permanent disability.

[8] We assume that benefits are computed based on full-time weekly earnings. This assumption reflects the law in many states, and the actual work week for the majority of the sample.

sample rather than a high-risk sample. We will use the death risk variable as a proxy for the overall job risk, since this measure is available on a state-specific basis, whereas published nonfatal injury data are not readily available.[9]

The basic workers's compensation variable in Table 1 is the real dollar value of the state weekly workers' compensation maximum benefit, WCMAX, which averages \$207. Benefits are computed using information on the worker's state of residence, martial status, number of dependents, and the formula given above. We also compute a benefit variable following Moore and Viscusi (1989), using the maximum benefit level in each state as a proxy for expected benefits. Since changes in the maximum influence the distribution of benefits for which each worker can potentially qualify, this variable measures changes in expected benefits when there is wage rate uncertainty. To control for the fact that increases in the maximum are more highly valued by workers whose wage places them at or above the maximum, we estimate the effect of this variable separately for each class of worker. The relative sizes of the estimated effects provide a check on the plausibility of our results. These results also provide a check on the robustness of the results derived using the actual benefit level in the replacement rate.

As shown in Viscusi and Moore (1987), workers only value accident insurance at positive risk levels. Thus, the appropriate measure of workers' compensation, the weighted weekly benefit level, involves an interaction of the death risk and the benefit level. Insurance benefits are captured by the risk-weighted replacement rate. The benefit level for which the worker qualifies is interacted with the risk variable to create the weighted benefit measure and then divided by the worker's after-tax weekly wage to create the weighted replacement rate variable. This formulation recognizes the fact that the value of the benefit varies with the risk and the fact that workers' compensation benefits are tax exempt. Use of the weighted replacement rate variable is consistent with much of the previous research on the wage effects of workers' compensation, although some studies have entered wages and benefits separately.[10]

As Table 1 indicates, the average replacement rate for the workers in each tenure group is 0.70. This rate is slightly higher than the nominal replacement rate of 0.66 used by most states. Two competing influences lead to this divergence. Since our replacement rate is computed on an after-tax basis, it will tend to be higher than the before-tax nominal rate of 0.66. On the other hand, since many workers' wages put them above the maximum, the observed replacement rate will tend to be lower. The net effect is to yield a rate slightly above the state-mandated nominal rate.

Empirical Results

The empirical tests of the worker learning hypothesis compare the wage-benefit and wage-risk tradeoffs estimated in quit equations across two tenure groups—workers with at least three years of tenure, and those with less than three years. We focus on the quit equation rather than a wage equation because the primary matter of interest is how the wage-risk and wage-benefit preferences of different groups of workers are altered, not how market contracts respond. The cutoff point at three years of tenure is the division used in Kahn's (1987) analysis of the preferences of marginal workers. Furthermore, restricting the newly hired worker sample to two years of tenure or less, as in Viscusi (1980c), yielded very small samples.

Estimates of the parameters of the quit equations for the senior worker group provide information on the behavior of senior workers on the margin, whereas estimates of the parameters of the quit

[9] Published single-digit (SIC) injury rates for fatal and nonfatal injuries exhibit a correlation of about .70 for 1986, significant at the .10 level. See U.S. Department of Commerce (1989), Tables 681 and 682.

[10] See Worral and Butler (1985).

Table 1. Variable Definitions and Sample Characteristics.

Variable Name	*Variable Definition*	*Mean (Standard Deviation)*	
		Tenure ≥ 3	*Tenure < 3*
EDUCATION	Years of education	12.93 (2.61)	13.23 (2.35)
FEMALE	1 if worker is female, 0 otherwise	0.11 (0.31)	0.21 (0.41)
HEALTH LIMITATION	1 if worker reports the presence of a health impairment that limits the amount of work he can do, 0 otherwise	0.07 (0.26)	0.05 (0.22)
DEPENDENTS	Number of dependent children	1.05 (1.13)	0.97 (1.09)
SINGLE	1 if worker has never been married, 0 otherwise	0.09 (0.28)	0.20 (0.40)
EXPERIENCE	Years worked for pay since age 16	21.17 (12.15)	13.61 (10.08)
UNION STATUS	1 if worker's job is covered by a collective bargaining agreement, 0 otherwise	0.34 (0.48)	0.23 (0.42)
WAGE	Worker's after-tax hourly wage in 1982 dollars (GNP deflator)	7.11 (2.33)	6.31 (2.45)
RISK	NTOF fatality rate variable: number of fatalities per 100,000 workers, by state and one-digit (SIC) industry	7.54 (9.81)	8.43 (9.50)
WCMAX	Maximum benefit level for temporary disability in the worker's state	210.81 (69.89)	195.35 (65.50)
REPLACEMENT RATE	Portion of weekly after-tax wage replaced by workers' compensation	0.70 (0.20)	0.70 (0.18)
d	1 if 2/3 of worker's weekly wage exceeds WCMAX, 0 otherwise	0.74 (0.44)	0.66 (0.47)
STRONG QUIT INTENTIONS	1 if worker is looking very hard for a new job, 0 OTHERWISE	0.08 (0.27)	0.15 (0.36)
WEAK QUIT INTENTIONS	1 if worker is looking at least somewhat hard for a new job, 0 otherwise	0.14 (0.34)	0.22 (0.42)
ACTUAL QUITS	1 if worker changed jobs in the past year, 0 otherwise	0.03	0.10
SAMPLE SIZE		2007	564

equations for the junior worker group provide information on the preferences of the newly hired worker. Our principal hypothesis is that the wage-benefit tradeoffs for the newly hired worker group will be less negative than those of the marginal senior worker group, which will include many workers who have acquired unfavorable information about risks on the job. We expect the risk effect to be larger in the senior worker group for the same reason. The self-selection of workers with more extensive learning and more precise risk perceptions into the junior worker group will tend to work against our principal hypotheses.

The quit equations are estimated using three measures of quit behavior. Table 1 defines these variables. If a worker answers "yes" to a question asking whether he is considering looking for a new job, the weak quit intention variable equals one; if "no," it equals zero. A similarly constructed measure of strong quit intentions equals one if the worker reports that he is seriously considering a new job and zero otherwise. Finally, the actual quit variable equals one if the worker quit during the year and zero otherwise. This variable pertains only to the 1981 and 1982 data. The quit variables reflect aggregate quit behavior fairly closely, as the average quit rate in manufacturing industries equaled about 1.5% per month in the late 1970s.[11] The average value of the weak quit intention variable of 14%, or

[11] See U.S. Department of Labor (1977).

1.2% per month, roughly equals the observed rate. The actual quit rate in our sample, 4.5% per year, is lower than the aggregate manufacturing rate, as expected, given the broader mix of industries represented in our sample.

The quit equations estimated are of the form

$$\text{(13)} \quad \text{Quit}_i = [1 + \exp - (\alpha' X_i + \phi_q \text{Wage}_i + \gamma_q \text{Risk}_i + \delta_q \text{Weighted Replacement Rate}_i)]^{-1} + \epsilon_{iq}.$$

Due to the binary nature of the dependent variable in these equations and the presence of the endogenous wage and replacement rate variables on the right-hand side of equation 13, nonlinear two stage least squares is used to estimate the parameters of the model.[12] As shown by Amemiya (1985), these estimates are both consistent and asymptotically normal. Instrumental variables include all of the exogenous explanatory variables in the quit equations and state dummy variables.[13]

Higher worker wages should reduce quitting by increasing the attractiveness of the worker's current job. Quit rates should increase with risk levels if there are learning-induced quits, and higher workers' compensation benefits should diminish quitting. The coefficients of interest are ϕ_q, γ_q, and δ_q, which we will use to calculate the wage–workers' compensation and wage-risk tradeoffs implied by workers' quit behavior.

Table 2 presents estimates of equation 13, using the actual quit variable, for workers in each tenure group. The actual quit variable provides a strong measure of the job satisfaction of workers, since workers are less likely to quit than to merely seek a new job. This measure should consequently reflect most strongly the role of worker learning in affecting the wage-risk and the wage–workers' compensation tradeoffs.

The Table 2 results indicate that wages and job risk characteristics are the most important determinants of workers' quit behavior. For workers with more than three years of tenure, increases in the wage exert significant downward pressure

Table 2. Determinants of Quits: 857 Workers, 1981–1983.
(Standard Errors in Parentheses)

Independent Variable[a]	*Tenure ≥ 3*	*Tenure ≤ 3*
EDUCATION	0.366**	−0.525***
	(0.167)	(0.211)
FEMALE	−0.832	−0.549
	(0.795)	(0.659)
HEALTH LIMITATION	0.408	−0.591
	(0.641)	(1.632)
DEPENDENTS	0.418**	−0.306
	(0.227)	(0.293)
SINGLE	0.534	0.398
	(0.671)	(0.670)
EXPERIENCE	−0.038*	−0.073*
	(0.029)	(0.046)
UNION STATUS	−0.050	−0.361
	(0.568)	(0.557)
ln (Weekly Wage)	−4.111***	3.076**
	(1.430)	(1.472)
RISK	0.590***	−0.028
	(0.215)	(0.144)
WEIGHTED REPLACEMENT RATE	−1.673***	−0.141
	(0.659)	(0.251)
CONSTANT	18.029	−12.826
	(7.098)	(6.472)

Sources: Michigan Panel Study of Income Dynamics, 1981, 1982, 1983; National Institute for Occupational Safety and Health data for 1981–85.

[a] Also included as explanatory variables are a Southeast regional dummy variable, a city size variable, and a year dummy variable.

* Statistically significant at the .10 level: ** at the .05 level; *** at the .01 level (one-tailed tests).

[12] The SAS procedure SYSNLIN, with the Gauss-Newton Minimization method, is used to estimate the model.

[13] We experimented with the use of age, experience, and tenure variables as instruments. Since these variables are important predictors of the wage, they would serve as useful instruments if they were independent of the error term ϵ_{iq}. When experience and tenure variables are added to he vector of instruments, there is no change in point estimates of the coefficients. The estimated standard errors are larger when the age variable is used in the weak quit equation. The main results, particularly those in the actual quit equation, are unaffected. We report results using the age, experience, and tenure variables as instruments.

on quits. Similarly, increases in workers' compensation benefits (as captured by the weighted replacement variable) decrease quits by workers who have been on the job for three years or more. Furthermore, among senior workers, increases in the risk level have a significant positive effect on quit behavior, as on-the-job experience makes workers more aware of adverse working conditions and also increases the precision of their estimates of the probability of a job-related injury or health problem. This risk effect is the same as that found by Viscusi (1979a). The wage effect, too, mirrors the earnings effect in Viscusi's quit intention equation. The additional effect shown by the workers' compensation variable provides further support for the model of rational worker learning: as workers perceive their jobs to be more dangerous, workers' compensation benefits become more valuable to them and serve to dampen the influence of risk on turnover by mitigating the financial losses associated with an accident.

The remaining variables in the quit equation measure the effect of the worker's characteristics on quits, holding constant the wage, the job risk, and the weighted replacement rate. In most cases, the signs of these variables are theoretically indeterminate. Significant effects are found, however, for education, number of dependents, and work experience in the more senior worker group.

The equation for workers with less than three years of tenure indicates that roughly the same control variables exert a significant effect on worker quits as in the senior group. In one case, however, the effect is opposite that for the older group: education increases quits by the older workers but decreases quits by younger workers. This result could reflect the net effects of a number of influences. Educational attainment may affect one's prospects for external mobility in a manner that varies with the extent of one's job-specific experience. Education may also complement specific training, which is accumulated in the early years of a job. This complementarity would tend to reduce quits initially. The negative education effect for junior workers could also reflect firms' greater commitment to preserving the match for more educated young workers. As workers acquire more work history, the job-signaling information content of the education becomes less important to the employer. For senior workers, greater education increases job mobility, with little connection between a worker's education and the firm's desire to retain the worker.

Consistent with our theoretical predictions, none of the job risk characteristics variables are statistically significant in the equation for junior workers, because their perceptions of risks and, therefore, their valuations of risk insurance are very imprecise. Indeed, in the early stages of the employment process, workers have been shown by Viscusi (1979a) to show a systematic preference for jobs with characteristics that are only poorly understood. This preference is reflected in the lack of an effect of the risk variable and the weighted replacement rate variable.

The wage variable exerts a significant positive effect for the junior workers. This unexpected result could be due to a number of influences. The most obvious explanation, that the causality between quits and wages runs in the opposite direction, not only would have to survive the instrumentation but also would have to provide an explanation of why reverse causality only matters for junior workers. Alternatively, wages could be acting as a proxy for skills that junior workers are offering to different employers. These skills could be observable to both the worker and the firm, unlike those in the signaling context discussed above, but not captured in our data. More able workers and firms would be willing to invest more in the job matching process. Furthermore, in the job matching model of Mortensen (1978), wages are not necessarily negatively related to turnover. Rather, the wage acts as a proxy for the match-specific capital; and inclusion of these components of the capital as regressors will eliminate any wage effect. Finally, it could be the case that the expected search costs are lower for high-wage young workers than for those bound by minimum wages, who face large queues for available jobs, thus making the high-wage junior workers more likely to quit.

For the actual quit equation, we thus have reasonably precise estimates of the parameters used to calculate the wage–workers' compensation tradeoff for workers with more than three years of tenure. The wage, risk, and workers' compensation variables are all significant at the 1% level or lower for these workers. As expected, variation in benefits does not cause any significant variation in quits for new hires.

Further support for the learning model is found by comparing the effects of RISK on quit intentions and on actual quits across tenure groups for alternative specifications. Table 3 presents the risk and workers' compensation effects for all three measures of quit behavior, as well as the effects of different measures of the wage variable (the after-tax weekly wage is always used in the replacement rate variable). In addition to the actual quit variable, we also estimate

Table 3. Wage-Risk and Wage-Replacement Rate Tradeoffs: Summary of Coefficient Estimates. (Standard Errors in Parentheses)

Dependent Variable		*Wage*	*Risk*	*Weighted Replacement Rate*
Actual Quits				
(i)	ln (Weekly Wage)			
	TENURE ≥ 3	−4.111***	0.590***	−1.673***
		(1.430)	(0.215)	(0.659)
	TENURE < 3	3.076**	−0.028	−0.141
		(1.472)	(0.144)	(0.251)
(ii)	Weekly Wage			
	TENURE ≥ 3	−0.012**	0.650***	−1.819***
		(0.005)	(0.245)	(0.690)
	TENURE < 3	0.004**	0.003	−0.130
		(0.002)	(0.115)	(0.221)
Weak Quit Intentions				
(i)	ln (Weekly Wage)			
	TENURE ≥ 3	−2.246***	0.077*	−0.122*
		(0.573)	(0.051)	(0.083)
	TENURE < 3	0.458	−0.031	0.042
		(0.681)	(0.057)	(0.076)
(ii)	Weekly Wage			
	TENURE ≥ 3	−0.005**	0.072*	−0.115*
		(0.002)	(0.048)	(0.079)
	TENURE < 3	0.04E-3	−0.009	0.014
		(1.58E-3)	(0.053)	(0.072)
Strong Quit Intentions				
(i)	ln (Weekly Wage)			
	TENURE ≥ 3	−2.104**	0.085**	−0.111**
		(0.646)	(0.052)	(0.087)
	TENURE < 3	3.077**	−0.028	−0.141
		(1.472)	(0.144)	(0.251)
(ii)	Weekly Wage			
	TENURE ≥ 3	−0.005**	0.087**	−0.112*
		(0.002)	(0.050)	(0.083)
	TENURE < 3	[a]	[a]	[a]

Sources: See notes to Table 2.
[a] Estimates would not converge.
* Statistically significant at the .10 level; ** at the .05 level; *** at the .01 level (one-tailed tests).

equation 13 using measures of weak and strong quit intentions as dependent variables. As the results in Table 3 indicate, the findings in Table 2 are quite robust over the different specifications. Both higher wages and increases in the weighted replacement rate significantly reduce quits and quit intentions for workers in the more senior tenure groups. For workers with less than three years' tenure, actual quits are positively and significantly related (as before) to wages; neither risk nor the weighted replacement rate shows any systematic effects on quits or quit tendencies for these workers.

To evaluate the robustness of our results, particularly with respect to the specifications used in some of our previous work, we also estimated the quit equations using the benefit measure of Moore and Viscusi (1989), which analyzed the effects of workers' compensation on job fatalities. As noted above, the numerator of the weighted replacement rate variable in this case uses the maximum benefit payment for the worker's state as the numerator of the replacement rate. To account for the fact that changes in the benefit maximum will be more highly valued by the workers whose wage exceeds the maximum, we estimate δ_q separately for each worker group (that is, for those workers whose current weekly wage places them above or below the maximum). This comparison is accomplished using the dummy variable *D*, defined earlier. The variable *D* is also treated as endogenous in estimating the quit equations. An additional test of the plausibility of our results is the prediction that the estimate of δ_q, given $D = 1$, should be more negative than its estimated value when $D = 0$.

Use of the benefit maximum as the numerator could introduce some error into our measure of the replacement rate, since it overstates the replacement rate for workers whose wages are low enough to place them below the maximum in their state. In Moore and Viscusi (1989), however, we argue that changes in the maximum will be valued by all workers because weekly wages are not known with certainty *a priori*. An increase in the maximum will therefore increase expected benefits for all workers, with the extent of the increase being felt most strongly by workers whose wage places them above the maximum. Finally, since benefit ceilings are one of the primary policy instruments available for altering benefit levels, direct estimates of the effect of changes in the maximum are most relevant for policy purposes.[14]

The results of this estimation, summarized in Table 4, essentially replicate those reported in Table 3 in terms of sign and statistical significance. The quit behavior of workers with more than three years of tenure is systematically related to both the risk level and the risk-weighted value of workers' compensation benefits. Increases in risk lead to significantly more quits, and to increased intentions to quit. This effect is once again reduced by insurance for financial losses and medical costs associated with an injury that are embodied in the workers' compensation program. Quits and quit intentions of workers with less than three years' tenure, on the other hand, are not significantly related to either of these forces.

A further test of the plausibility of the model compares the coefficients on the weighted replacement rate variables for workers whose wage places them above or below the benefit maximum. If workers are uncertain about their future wage, then each worker will attach some likelihood to the possibility that the wage at the time of an injury will exceed the benefit maximum. Workers whose wages currently exceed the maximum will attach a greater probability to this outcome and will, therefore, value changes in the benefit maximum more highly. As a consequence, the estimated replacement rate effects should be more negative for workers whose current weekly wage exceeds the benefit maximum.

Our results support this prediction for the

[14] A further reason for use of the maximum is that it allows direct comparisons with our other published studies on this subject (Moore and Viscusi 1990). An earlier version of this paper used the benefit maxima rather than the replacement rate as the benefit measure. The signs, significance levels, and magnitudes of the key coefficients in the wage and quit equations mirror those reported here.

Table 4. Wage-Risk and Wage-Benefit Tradeoffs: Summary of Coefficient Estimates Using Alternative Benefit Variable.
(Standard Errors in Parentheses)

Dependent Variable	*ln (Wage)*	*Risk*	*Weighted Replacement Rate* × *d*	*Weighted Replacement Rate* × *(1-d)*
Weak Quit Intentions				
TENURE ≥ 3	−2.022***	0.266**	−0.579**	−0.205**
	(0.680)	(0.115)	(0.290)	(0.101)
TENURE < 3	1.193	−0.062	0.091	−0.101
	(1.025)	(0.087)	(0.124)	(0.094)
Strong Quit Intentions				
TENURE ≥ 3	−2.200***	0.411***	−1.028**	−0.290**
	(0.750)	(0.159)	(0.464)	(0.139)
TENURE < 3	−0.422	0.042	−0.129	−0.026
	(0.904)	(0.086)	(0.176)	(0.072)
Actual Quits[a]				
TENURE ≥ 3	−0.070**	5.73E-3***	−8.45E-3***	a
	(0.042)	(2.32E-3)	(2.94E-3)	
TENURE < 3	0.151	−2.28E-3	−3.09E-3	a
	(0.106)	(5.58E-3)	(6.82E-3)	

Sources: See notes to Table 2.
[a] Not significantly different from adjacent estimates.
* Statistically significant at the .10 level: ** at the .05 level; *** at the .01 level (one-tailed tests).

senior worker group in two of three cases. For both quit intention variables, the weighted replacement rate effects are negative and significant whether the wage exceeds or falls below the maximum. Furthermore, the estimated coefficient is always more negative than the corresponding estimate for low-wage workers. In the actual quit equation estimated for more senior workers, the two estimates are not significantly different from each other. When restricted to equality, they are negative and significant. Consistent with our principal hypothesis, the estimated wage-benefit tradeoffs are less negative for the junior worker group. This pattern obtains for both low- and high-wage workers, regardless of the quit variable used. These results, and the general robustness of the results in Table 3, indicate that the hypothesis withstands a variety of changes in specification.

Tests of Worker Learning–Wage Benefit Tradeoffs

The model predicts that the wage-benefit tradeoff should be more negative for senior workers on the quit–no quit margin than for the new hires. In an important sense, the detailed calculations are unnecessary, since none of the tradeoffs for the junior worker groups are based on coefficients that differ significantly from zero. Nonetheless, point estimates will indicate whether the hypotheses hold for the data in our sample. Using the results in Table 3, the tradeoffs can be calculated directly.

In the quit equations, the wage-benefit tradeoff is computed as the negative of the ratio of the partial effect of a dollar increase in workers' compensation benefits on quit intentions, $-(\partial Q/\partial b)$, to the effect of a dollar increase in wages on the same dependent variables, $(\partial Q/\partial w)$, where b and w denote the benefit and the wage:

$$\frac{\partial w}{\partial b} = \frac{-\partial Q/\partial b}{\partial Q/\partial w}. \tag{14}$$

To evaluate this expression, the quit equation given by equation 13 above is rewritten as

$$(15)\quad Q = [1 + \exp - (\alpha' X + \phi_q wH + \gamma_q p + \delta_q pR)],$$

where w denotes the wage, p the death risk, H the hours worked per week, and R the replacement rate.[15] Letting P(Q) denote the right-hand side of this equation, the partial effect of an increase in the wage on quits and quit intentions equals

$$(16)\quad \frac{\partial Q}{\partial w} = \mathrm{P}(Q)\ (1 - \mathrm{P}(Q))\ (\phi_q H + \delta_q p \frac{\partial R}{\partial w}).$$

The replacement rate can be written as a function of the weekly wage, the dummy variable D, the benefit ceiling WCMAX, and the tax rate:

$$(17)\quad R = \frac{(2/3)\ (1 - D)wH + D\text{WCMAX}}{w(1 - t)H}.$$

Differentiating this expression with respect to the wage then yields the expression

$$(18)\quad \frac{\partial R}{\partial w} = \frac{[w[\text{WCMAX} - (2/3)wH]\ \frac{\partial D}{\partial w} - D\text{WCMAX}]}{w^2(1 - t)H}.$$

The first term in this expression will always equal zero, since variation in the weekly wage will affect the variable D only when the worker just qualifies for the maximum, at which point the first part of this term equals zero. Inserting the nonzero portion of equation 18 into equation 16 yields the expression

$$(19)\quad \frac{\partial Q}{\partial w} = \mathrm{P}(Q)\ (1 - \mathrm{P}(Q))\ \frac{\phi_q H - \delta_q pD\text{WCMAX}}{w^2(1 - t)H}.$$

Again using equation 15 and the definition of the replacement rate $R = b/w(1 - t)H$, the effect of an increase in benefits on quit behavior equals

$$(20)\quad \frac{\partial Q}{\partial b} = \mathrm{P}(Q)\ (1 - \mathrm{P}(Q))\ (\delta_q p / w(1 - t)H).$$

The wage-benefit tradeoff then equals the negative of the ratio of equation 20 to equation 19,

$$(21)\quad \frac{\partial w}{\partial b} = \frac{-\delta_q p/(w(1 - t)H)}{\phi_q H - \delta_q pD\text{WCMAX}/(w^2(1 - t)H)}.$$

Similar calculations yield the expression for the wage-risk tradeoff

$$(22)\quad \frac{\partial w}{\partial p} = \frac{\gamma_q - \delta_q R}{\phi_q H - \delta_q pD\text{WCMAX}/(w^2(1 - t)H)}.$$

Table 5 summarizes the median wage-benefit tradeoffs, and also the median wage-risk tradeoffs computed using the coefficient estimates in Table 3. We report medians to control for outlier problems associates with computing the sample means.

In every case considered, as predicted by the theory, the median wage-benefit tradeoffs are more negative (and less than zero) for the senior workers. For the median senior worker in the quit equation, the wage-benefit tradeoff equals

Table 5. Wage-Benefit and Wage-Risk Tradeoffs: Median Estimated Effects.

Quit and Wage Variables	TENURE ≥ *3 years*	TENURE < *3 years*
Wage-Benefit Tradeoffs		
Actual Quits		
ln (Weekly Wage) (× 10^{-3})	−3.5	5.3
Weekly Wage (× 10^{-3})	−3.6	12.0
Weak Quit Intentions		
ln (Weekly Wage) (× 10^{-3})	−6.6	−0.2
Weekly Wage (× 10^{-3})	−7.6	48.9
Strong Quit Intentions		
ln (Weekly Wage) (× 10^{-3})	−6.4	5.3
Weekly Wage (× 10^{-3})	−7.4	a
Wage-Risk Tradeoffs		
Actual Quits		
ln (Weekly Wage) (× 10^{-2})	22.5	25.5
Weekly Wage (× 10^{-2})	12.2	37.9
Weak Quit Intentions		
ln (Weekly Wage) (× 10^{-2})	−4.4	−2.6
Weekly Wage (× 10^{-2})	−6.3	−20.8
Strong Quit Intentions		
ln (Weekly Wage) (× 10^{-2})	0.4	25.5
Weekly Wage (× 10^{-2})	1.7	a

Sources: See notes to Table 2.

[15] Note that we now use p to denote the risk of an injury, whereas the theoretical model used p to denote the probability of no injury.

−.35 cents per dollar of benefits, compared to a positive wage-benefit tradeoff of .53 cents per dollar of benefits for junior workers. In addition to the absolute differences observed in the quit equations and the similar differences in the weak and strong quit intentions equations, the wage-benefit tradeoffs for the senior workers are all based on coefficient estimates that are statistically significant, whereas those tradeoffs estimated for the junior workers represent a combination of coefficients, none of which are significantly different from zero. Finally, the wage-benefit tradeoffs for the senior workers are quite stable across all six equations. Thus, in addition to the observed differences in the tradeoffs between the more and less senior workers, the tradeoffs estimated for the senior worker group are also much more precise. We conclude that the predictions of the model are supported and appear to be robust across a variety of specifications of the dependent variables.

The wage-risk tradeoff results summarized in the second panel of Table 5 are less consistent with the learning model in which experienced workers have higher risk assessments. If workers were fully informed and all workers' perceptions were identical, the wage-risk tradeoffs would not vary by tenure group. In almost every instance, however, there is evidence of a larger positive wage premium for risk in the case of less experienced workers on the quit margin. Once again, however, few of the coefficient estimates in the junior worker group are significant, so this result is not particularly troublesome. Moreover, since the risk variable enters the equation directly and through the expected replacement rate variable, it may be difficult to reliably estimate the effects of these correlated variables.

Due to the insignificance of the coefficients in the junior worker equations, the most important information to be gained from these calculations lies not in the comparison of effects, but in the policy implications of the estimates for the senior worker group. Most important, the results support the view of the labor market as an efficient sorter of workers that compensates workers for exposure to risk and subsidizes injury insurance through wage reductions.

We can also use these results to estimate whether benefit levels are adequate, using the procedure in Viscusi and Moore (1987). In this framework, the observed tradeoff will equal the ratio $-p/(1-p)$, where p is the probability of an accident. For the fatality risk data, this ratio equals about $-1/10{,}000$. The estimated tradeoffs between benefits and wages in the actual quit equations, when both wages and benefits are put on a weekly basis, are many times greater than this figure.[16] Thus, since the observed tradeoff exceeds the optimal tradeoff, it would appear that benefit levels are too low. Since we do not include a nonfatal risk variable in the quit equations, however, the observed rate of tradeoff might reflect instead the rate that is optimal for risks of all types of injuries, which would be much closer to the observed tradeoff.

The same coefficients can also be used to calculate the value of life and the degree of risk aversion implied by our estimates. Using the estimates in the strong quit equation, the estimated value of life implied by the estimates ranges between 1 and 4.25 million dollars.[17] This range is lower than the estimates in the wage equations estimated by Moore and Viscusi (1989) using the NTOF data, and is similar to many estimates found in the literature. Since the wage-risk tradeoffs are not as stable as the wage-benefit tradeoffs across equations within the senior worker group, the value-of-life calculations are not as meaningful as the calculations for assessing benefit optimality.

Finally, since we know that workers will give up approximately .004 dollars in weekly wages for a one-dollar increase in benefits, we can compute the degree of

[16] The estimated tradeoff of .0035 dollars per hour multiplied by 40 hours yields a weekly wage decrease of 14 cents corresponding to a \$1 benefit increase.

[17] For example, the wage-risk tradeoff of .017, when multiplied by 2,000 hours, an inflation factor of 1.25, and 100,000 (the risk scaling factor that yields one statistical life), equals 4.25 million.

risk aversion exhibited by the workers in our sample. If benefits were to rise by two dollars per week, wages would fall by about \$.01, for an annual decline of 50 cents. The discounted expected value of this two-dollar benefit increase, at an annual risk level of 1/10,000 and with 52 weeks per year and a real discount rate of 5%, is 21 cents. Thus, the implication is that workers are risk averse, and are willing to sacrifice one dollar of wages for about 42 cents in expected insurance benefits.

Conclusion

The results in this paper and in other recent studies suggest that compensating differentials for risk should be viewed as a broader issue than the standard wage-risk tradeoff literature implies. Wages and workers' compensation serve as complementary compensation mechanisms, with wages providing ex ante risk compensation and workers' compensation providing ex post earnings replacement. Each of these earnings components reduces worker quitting, and workers accept a lower wage in return for higher workers' compensation benefits. This tradeoff reflects worker preferences for different forms of risk compensation.

The results presented here utilize these wage and quit relationships to explore the differences in the wage–workers' compensation tradeoff for workers on the quit margin. Estimated tradeoffs indicate that workers on the quit margin place a much higher relative value on workers' compensation than do new hires. Moreover, greater job risks lead to significant increases in quits and quit intentions for workers on the quit margin but have no effect on the quit behavior of junior workers. These findings are consistent with a model in which worker quits are induced in part by learning about risks on the job.

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