Chapter 8

ENVIRONMENTAL REGULATION AND THE VALUATION OF LIFE: INTERINDUSTRY MOBILITY AND THE MARKET PRICE OF SAFETY*

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1 INTRODUCTION

Environmental policy in both Europe and the United States has increasingly become associated with the necessity to value costs and benefits of proposed regulations. This growing emphasis on "cost-benefit" analysis is particularly acute for health and safety hazards (and regulations) that are difficult to measure and involve relatively small probabilities. In these instances the fundamental question is how to measure the benefits of a regulation that saves a human life. As indicated in Table 1, current methods of assigning a dollar value to human life by government agencies in the United States indicate a wide variance. The use of these estimates of human life in setting environmental regulations can be controversial. As discussed by Waldman (1988), in a 1986 suit against OSHA the Public Citizen Litigation Group charged that cost-benefit analysis was used to weaken regulation of ethylene oxide, a sterilizing agent that may cause cancer and spontaneous abortions. Similarly, the continuing controversy over EPA draft rules which appeared initially in 1984 phasing in a ban on asbestos has often centered on the proper discounted value of life. These examples clearly indicate that environmental policy and the valuation of life are becoming increasingly intertwined.

Economists have often approached the issue of determining the value of life from the perspective of compensating wage differentials. The theory of compensating wage differentials, attributable to Adam Smith, suggests that jobs with disagreeable characteristics will command higher wages, ceteris paribus. Most empirical tests of this theory with hedonic wage equations have addressed the job-related risks of death or serious injury,

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TABLE 1

Value of Life Estimates: U.S. Government Agencies.

Agency	Value of Life ($)
Consumer Product Safety Commission (CPSC)	2 million
Environmental Protection Agency (EPA)	475,000 - 8.3 million
Federal Aviation Administration (FAA)	1 million
Occupational Safety and Health Administration (OSHA)	2 million - 5 million

Source: Various government publications.

and, for the most part, have found such compensation to indeed exist. For examples of U. S. experience, see Dillingham (1985), Dorsey (1983), and Gegax, Gerking and Schulze (1986), while risk compensation in Britain and Sweden has been examined by Martin and Psacharopoulos (1982) and Duncan and Holmlund (1983), respectively. Implicit within this literature is the assumption that workers' willingness to pay for risk reduction (safety) in the workplace through diminished wages, and market valuations of the "price" of these reductions, are equivalent. Thus, it is generally assumed that no externalities exist in the pricing and provision of industrial safety, and therefore that estimates of the market price of incremental safety also measure the wage-risk trade-off of workers. Our analysis below suggests that this is clearly not the case within the manufacturing sector, where willingness to pay for risk reduction exceeds the price (cost) of such reduction when measured at current levels of risk exposure. Thus, to the extent that implicit prices derived from (hedonic) wage equations understate worker willingness to pay for incremental safety, implied value of life estimates determined from these equations are downward biased.

In the following section, risk valuation is examined within the context of compensating wage differentials. A demand-side method for valuing willingness to pay for risk reduction is also proposed, and compared to that employed to generate market prices for enhanced safety in the workplace. These two methods, or models, for valuing the wage-risk trade-off are specified in Section 3 by examining their individual determinants and sources of data. Econometric estimates of these models are then presented and discussed in Section 4, and provide the basis for a comparison of the implied value of life evaluated, alternatively, by market price and willingness to pay. Conclusions follow in Section 5.

2 THE VALUE OF LIFE AND RISK VALUATION

2.1 Measuring the Wage-Risk Trade Off

Jobs can be characterized by many dimensions such as likelihood of injury, pace of work, and the general unpleasantness of tasks to be performed. Employees and employers negotiate a single wage at the time of hiring that reflects both underlying demand and supply functions pertinent to these and other job characteristics. Under the assumption of perfectly competitive labor markets (and thus perfect information on working conditions), an equilibrium price is established for each job attribute that is equal to its marginal cost. Such "compensating differentials" among wages for various jobs are often examined, and empirically estimated, based upon the hedonic theory of prices as outlined by Rosen (1974).

Following Rosen, consider jobs in which equilibrium quantity and value, or market-clearing implicit price, have been established for all job attributes but one, here the risk of injury, R. In order to maintain a constant level of profit, firms will supply safer jobs only at reduced wages, W, ceteris paribus.[1] Again assuming that equilibrium value and quantity have been predetermined for all job attributes other than risk, worker indifference curves representing various levels of total utility and an associated wage-risk trade-off can be determined. In this context, with worker satisfaction held constant, the required trade-off between wages and risk increases with the level of risk exposure. Following Rosen (1974), a "market clearing implicit price" curve can be determined, W(R).

Within perfectly competitive labor markets, workers maximize utility by accepting jobs along W(R) that equate, at the margin, their willingness to "pay" for risk reduction (safety) with the wage-risk trade-off required by firms to maintain (zero) profits. Because workers differ in their tastes for job safety, various wage-risk bundles (W,R) are selected at long-run equilibrium. Thus, as noted by Smith (1979), "workers who value safety highly tend to accept jobs with firms that can offer it most cheaply." On the other hand, a worker who places a lower value on safety will seek employment with a firm offering both high relative wages and risk. The coefficient estimate on risk (or its transformation) in a hedonic wage equation can be interpreted as the slope of W(R) at the prevailing level of risk exposure, W'(R). In addition, within perfectly competitive

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[1] For a detailed discussion of this theory see Rosen (1974). Also, see Curington (1986) on the impact of OSHA safety regulations on both workplace injuries and their frequency.

labor markets in long-run equilibrium, this slope is also equal to a representative worker's willingness to pay for risk reduction.

However, there are several reasons to suspect that labor markets are not in a state of equilibrium, and therefore, that estimates of W'(R) obtained from (hedonic) wage equations do not approximate the representative worker's offered wage concession for added safety at the prevailing level of risk exposure in the economy. Conditions existing within imperfect labor markets that likely maintain this imbalance include imperfect information, ineffective bargaining, and transactions costs.

With respect to imperfect information, Viscusi (1979) and Viscusi and O'Connor (1984) found worker quit rates to be positively augmented by job hazards; and that the imperfect nature of workers' prior information combines with the potential for learning on the job to produce this adaptive behavior. On the other hand, ineffective collective bargaining, or any noncompetitive aspect of the contracting process, could also produce this situation. Such a case would arise if the risk component of a negotiated wage package fails to align firm and workers' wage-risk trade-offs (by a single "price"). Evidence of ineffective labor contracting is provided by Martin and Psacharopoulos (1982) who show that collective bargaining in Great Britain actually weakens compensating differentials obtained between more and less dangerous jobs. Finally, to the extent that transactions costs of a wage-risk adjustment exceed the benefits of such an adjustment, this discrepancy will also be maintained. (Weinberg, Friedman and Mayo (1981) have analyzed similar transactions costs incident to housing search under disequilibrium conditions.)

Thus, there is reason to suspect that the willingness to pay for risk reduction on the part of workers diverges somewhat from its market "price." The direction of this divergence is, of course, less certain, but is amenable to empirical investigation below.

Based upon the above, estimates of market price and/or willingness to pay (and thus any divergence) cannot be obtained from a hedonic wage equation whose price function for job attributes such as risk describes an equilibrium relationship (and thus equality of market price and willingness to pay). In addition, both Epple (1987) and Bartik (1987) have recently shown that equilibrium conditions of the hedonic model impose surprising restrictions on the error terms of such models, restrictions that rule out many seemingly natural estimation strategies. An estimate of the implicit market price of risk at the prevailing level of risk exposure [defined above as W'(R)] will be obtained as the coefficient estimate on risk (or its transformation) in a wage/earnings equation with regressors selected in accordance with Mincer's model of schooling, experience and earnings [Mincer (1974)]. Such an equation, unconstrained by hedonic equilibrium conditions, can be represented as

$$W = W(H, X, R), \tag{1}$$

where W again represents wage, H and X are vectors of human capital and other controlling variables, respectively, and R is a measure of risk in the work place. Included in X are working conditions other than risk exposure.

On the other hand, a demand-side, or willingness to pay, interpretation of risk evaluation must measure the slope of a representative worker's indifference curve at the present level of risk exposure. However, since worker utility can not be measured, how is one to evaluate the worker's offered concession for added safety, dW/dR, while holding such worker satisfaction constant?

Under the assumption that workers in effect "vote with their feet" among jobs in various industries in response to working conditions and compensation received, then worker satisfaction (utility) is likely maintained to the degree that interindustry mobility is also maintained.[2] This mobility, here termed the likelihood of industry switching, S, can be represented as

$$S = S(H, X, R, W^*), \tag{2}$$

where the vector of human capital variables, H, controls not only for mobility variation directly attributable to factors such as education and experience, but also for the "cost" of job-search as well.[3] When combined with equation (1) to generate a first-stage estimate of W [W* in equation (2)], and holding the likelihood of industry switching constant (dS = 0), note that

$$\frac{dW}{dR} = -\frac{S'(R)}{S'(W)}, \tag{3}$$

by the implicit-function rule. To the extent that this willingness to pay for risk reduction departs significantly from W'(R), the provision of safety in the work place diverges from its "optimal" level.

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2 Tiebout (1956), of course, coined this phrase in describing household response to local government expenditures.

3 For detailed analyses of interindustry mobility and its determinants, see Gallaway (1969), and Schlottmann and Herzog (1984).

3 MODELS OF WAGE DETERMINATION AND INDUSTRY SWITCHING

3.1 Determinants

As will be explained below, equations (1) and (2) were fit to Census microdata for white-male workers employed within manufacturing. Personal characteristics entered within these equations to represent human capital include both age and education, as well as their interaction and squared terms.[4] Vocational training and work disabilities were also entered. To the extent that such "inputs" build human capital, wages should be augmented accordingly. Other personal characteristics of workers were included in both the wage and industry switching equations as control variables [X in equations (1) and (2)]. These include two measures of family dependency, namely marriage and the presence of school age children, as well as variables representing specific occupations. In addition, a variable representing prior migration was included in the industry switching equation to partially control for past interindustry mobility (which should augment present mobility).[5]

Two other groups of variables were included as determinants of both wages and industry switching [and comprise X in equations (1) and (2)]. The first of these relate to the labor market in which each worker resides. Economic conditions there are represented by the local unemployment rate and growth rate of total employment, while population and its density represent unmeasurable aspects of the labor market to include job-search (transactions) and living costs as well as amenities and disamenities. It is expected that nominal wages will be augmented, ceteris paribus, by tight labor markets and higher costs of living (population and population density). In addition, the likelihood of industry switching should be increased somewhat within labor markets characterized by rapid job creation.

Industry characteristics comprise the final group of variables relevant to the analysis. Based upon the industry of employment for each worker and data for those industries at the national level, variables representing employment growth, percent union, and the risk of fatal injury [the risk variable, R, in equations (1), (2) and (3)] were included within both the wage and industry switching equations. Each of these three variables is expected to augment worker compensation, while the likelihood of industry switching

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[4] These determinants were selected in accordance with Mincer's model of schooling, experience and earnings [Mincer (1974)]. In addition, see Blinder (1976).

[5] For analyses of these interactions, see Schlottmann and Herzog (1984).

should be diminished and increased respectively by industry growth and risk exposure. In addition, an interaction term set equal to the product of the risk and percent union variables was included within the wage equation to examine the effect, if any, of collective bargaining on risk rewards in manufacturing. Finally, based upon equation (2), each worker's wage [predicted in equation (1)] was considered a determinant of industry switching.

3.2 Data

Observations for the estimation of equations (1) and (2) comprise micro-data on personal characteristics (to include wages and industry switching) matched to aggregate data on both labor market of residence and industry of employment. As stated, workers were required to be white-males employed within manufacturing. In addition, workers were required to hold jobs as craftsmen, operatives or laborers in order to place them "at risk" to the job hazard variable. Table 2 presents the distribution of these job categories and representative occupations within each category. Based upon these and other restrictions, individuals were drawn from the 5% one-in-a-thousand Public Use Sample (PUS) of the 1970 Census (1972) in two nonexclusive groups: Group A representing workers employed within manufacturing in 1965, and Group B representing equivalent employment in 1970.[6] Observations drawn from the PUS were 4,511 in Group A and 4,509 in Group B. Finally, Group A (B) observations were matched in labor market and industry characteristics based upon residence and industry of employment in 1965 (1970), respectively. These two samples were necessary in order to fully exploit the wage information (see estimation below).

Labor market characteristics for states were obtained from published Cen sus materials and, with the exception of employment growth, take values for 1965 (1970) when matched

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[6] Workers were also required to be the chief income recipient within the family, and of age 19-55 in 1965 (24-60 in 1970). In addition, workers were deleted from the sample if they attended college in 1965 and/or 1970, or were members of the armed forces in 1965 and/or 1970. Finally, 1970 rather than 1980 Census microdata was employed in the study since the latter provides no coding to identify interindustry mobility.

TABLE 2

Sample Distribution of Manufacturing Employment and Representative Occupations.

Category	Percent of Sample[a]
Craftsmen and Kindred Workers:	40.5
Carpenters	
Cranemen, Derrickmen, and Hoistmen	
Electricians	
Heavy Equipment Mechanics	
Metal Job and Die Setters	
Operatives, Total:	52.3
(a) Operatives (Excluding Transport)	46.8
Assemblers	
Cutting Operatives	
Textile Operatives	
Machine Operatives	
Welders	
(b) Transport Operatives	5.5
Truck Drivers	
Forklift and Tow Motor Operatives	
Motormen in a Mine, Factory, etc.:	
Railroad Switchmen	
Bus Drivers	
Laborers	7.2
Construction Laborers	
Carpenter's Helpers	
Longshoremen	
Warehousmen	
Stockhandlers	

[a] 6,652 observations, Public Use Sample of the 1970 Census.

to Group A (B) observations. Among the industry characteristics, both employment growth and the percent union were derived from U. S. Department of Labor (1975), while the risk variable was determined on the basis of information provided in U. S. Department of Labor (1971).

In this study, workplace risk is represented by the number of on-the-job fatal injuries per million hours worked in specific three-digit Standard Industrial Classification (SIC) manufacturing industries.[7] Means of this "risk" in 1969 within the Group A and B samples are .0498 and .0490 respectively, and range from .0019 in electronic computing equipment to .4224 in logging.

4 ECONOMETRIC ESTIMATES

4.1 Wage (Earnings) Determination

Estimates of the wage equation were obtained for both W and ln(W) in equation (1), employing alternatively Group A and B data in order to examine the robustness of implied market prices of risk reduction. In all cases, the dependent variable was represented by weekly wage and salary earnings (or their natural logarithm) rather than by an hourly wage.[8] In addition, binary independent variables were created and set equal to unity (vs. zero) for married individuals as well as those with school age children, vocational training and/ or a disability which limits work. Also, specific manufacturing occupations were represented by dummy variables (with laborers excluded).

Ordinary least squares (OLS) estimates of individual 1969 weekly earnings are provided in Table 3 for white-male workers employed within manufacturing in 1970 (Group B data) and, in addition, corresponding estimates for 1965 (Group A data). In order to facilitate our analysis, we will confine our discussion to the 1969 estimates. Coefficient estimates are shown for both earnings and the natural logarithm of earnings, the latter form implying a rising price (reward) per unit of fatal injury risk. Notice that, with few exceptions, estimates in both equations satisfy a priori expectations.

Employing the semilogarithmic form as an example, note in Table 3 among human capital variables that both age and education significantly augment weekly earnings

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[7] Based upon U. S. Department of Labor (1971), this risk variable was determined as the product of the number of fatal or disabling injuries per million hours worked and the fraction of these injuries resulting in death. Assuming average weekly hours and weeks worked per year of 40 and 50 respectively, the annual likelihood (per worker) of a fatal injury is equal to the above number divided by 500. Information on "perceived" job related accidental death risks has recently been developed by Gegax, Gerking and Schulze (1986).

[8] W in equation (1) may be obtained for individual workers in 1969 from our microdata samples by dividing annual earnings by the product of weeks and hours worked (in a year and week respectively). However, this is not advisable since both earnings and weeks worked were measured in 1969 while hours pertain to the Census reference week in the <u>following</u> year, i.e. 1970.

TABLE 3

Determinants of Weekly Earnings in Manufacturing: Ordinary Least Squares Estimates for White-Males[a].

Independent Variables:	Dependent Variables (1969): Earnings	Dependent Variables (1969): Ln(Earnings)	Dependent Variables (1965): Earnings	Dependent Variables (1965): Ln(Earnings)
Constant	4.3706	3.4452**	57.6843***	4.16***
Personal Characteristics:				
Age	3.1183***	.0296***	.4052	.0044*
Age squared	-.0375***	-.0003***	.0023	.0001
Education	-2.8108	.0609***	2.2912***	.0219***
Education squared	.2840***	-.0010*	.0572**	.0015**
Age x education	.0869**	-.0001	.0133	.0001
Vocational training	12.4227***	.0693***	1.9564***	.0181***
Disability which limits work	-10.2836***	-.0886***	1.6392*	.0160**
Married	12.1252***	.0773***	.6448	.0078
School age children	9.2931***	.0414***	.4235	.0037
Occupation:				
Craftsman or kindred worker	35.0363***	.2374***	4.9327***	.0426***
Operative (except transport)	10.1065**	.0962***	3.4106***	.0291***
Transport equipment operative	16.7834***	.1416***	2.7209*	.0174
Labor Market Characteristics:[b]				
Employment growth, 1965-1970 (%)	1.9056	.0083	.5597	.0005*
Unemployment rate (%)	-4.7798***	-.0331***	1.8102***	.0175***
Population (10^6)	1.9376***	.0116***	.0011***	.0001*
Population density (10^3)	12.1522***	.0603**	.0100	.0003
Industry Characteristics:[c]				
Employment growth, (1965-1970 (%)	.6730***	.0054***	.8403***	.0095***
Percent union	.4450***	.0032***	.5713***	.0054***
Fatal injury risk	175.9491**	1.6521***	178.5870***	1.9557***
Fatal injury risk x percent union	-2.8983*	-.0293***	2.9809***	.0311***
F-statistic	55.20	55.48	79.15	84.16
R^2 (adjusted)	.20	.20	.26	.28
Number of observations	4,509	4,509	4,510	4,510

* t-test significant at the 0.10 level.
** t-test significant at the 0.05 level.
*** t-test significant at the 0.01 level.

[a]All variables are defined in the text. Estimates were obtained using Group A and Group B microdata also described in the text.
[b]Based upon state of residence in 1970. Variables represent 1970 values except where noted.
[c]Based upon industry of employment in 1970 and national data. Percent union and fatal injury risk represent 1968 and 1969 values, respectively.

(albeit at a declining rate), and that such augmentation is also attributable to vocational training. However, when work effort within manufacturing is limited by a disability, earnings suffer. On the other hand, such effort is apparently enhanced (resulting in increased earnings) among white-male workers encumbered by family responsibility. In addition, craftsmen and kindred workers as well as operatives (both transport equipment and other) receive compensation in excess of that awarded laborers, ceteris paribus.

Turning now to labor market characteristics in Table 3, note that manufacturing workers receive increased nominal compensation when employed within tight labor markets (low unemployment rates) as well as in those markets characterized by higher costs of living (high population and population density). Also, notice among industry characteristics that rapid industry expansion (employment growth), greater bargaining power of workers (percent union), and increased risk in the workplace each augment weekly earnings.

Thus, compensating wage (earnings) differentials do exist within manufacturing as a reward for risk exposure. In addition, the coefficient estimates on the fatal injury risk variable in Table 3 indicate the implicit market price of additional safety at prevailing levels of risk exposure. Finally, note the negative and significant interaction term between risk and percent union in Table 3.[9]. Two additional equations employing Group A

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9 This result is consistent with findings by Martin and Psacharopoulos (1982) for collective bargaining in Great Britain. They suggest two reasons for this result. First, collective bargaining often takes place for broader groups than the occupations used here, and broad wage settlements could reduce the sensitivity of wages to risk. The second possible reason is that to the extent unions often press directly for safety improvements rather than using risk solely in wage bargaining, risk would play less of a role in earnings determination by collective bargaining. On the other hand, Thaler and Rosen (1976) find a positive interaction for the United States by employing a binary variable representing union affiliation. In commenting on their work, Kosters (1976) suggests that unions may be better equipped than nonunion workers to assemble reliable information on risk, and to utilize it effectively during the bargaining process.

data and variables equivalent to those in Table 3 were estimated utilizing weekly earnings in 1965 [for use as first-stage estimates of W* in equation (2)].[10]

Although not shown in Table 3, coefficient estimates and significance levels are similar to those obtained for 1969. For instance, for both dependent variables (1965 weekly earnings and the natural logarithm of these earnings), signs and significance levels on industry characteristics match those in Table 3. Thus, compensating wage differentials for risk exposure within manufacturing were also detected in 1965, as was the compression of such risk rewards by increased unionization.

4.2 <u>Interindustry Mobility</u>

As stated, an estimate of workers' willingness to pay for risk reduction can be obtained from an industry switching equation where the likelihood of interindustry mobility is determined, in part, by risk exposure on-the-job as well as compensation received. In this respect, equation (2) was determined from Group A microdata for white-male workers employed within manufacturing in 1965 and, thus, "at-risk" to interindustry mobility over the ensuing five-year period.

For each of the 4,511 observations, the dependent variable in this equation was set equal to unity (vs. zero) if a worker was employed in an industry in 1970 (manufacturing or nonmanufacturing) other than that in which employed in 1965. Based upon three-digit SIC codes, 994 manufacturing workers (22 per cent) switched industries between 1965 and 1970, some migrating interstate in the process. Binary logit estimates of the determinants of this interindustry mobility among white-males are provided in Table 4.

Based upon asymptotic t-values in the last column of Table 4, note among personal characteristics that the likelihood of industry switching is augmented by vocational training, by a disability which limits work, and by prior geographic mobility.[11] On the other hand, this likelihood is diminished somewhat by increased age (an effect that also

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[10] However, labor market variables (representing residence in 1965) assumed 1965 rather than 1970 values. In addition, both 1965 weekly earnings and industry characteristics were based upon industry of employment in 1965.

[11] Binary independent variables are equivalent to those employed for earnings determination in Table 3. The variable representing prior migration (as of 1965) was set equal to unity (and zero otherwise) if an individual resided in a state other than his birth in 1965.

TABLE 4

Determinants of 1965-1970 Industry Switching For Individuals Employed in Manufacturing in 1965: Binary Logit Estimates for White-Males[a].

Constant and Independent Variables:	Coefficient	Asymptotic t-value
Constant	5.4349	2.93***
Personal Characteristics:		
Age	-.1825	-4.11***
Age squared	.0015	3.46***
Education	.0727	.58
Education squared	-.0020	-.46
Age x education	-.0004	-.23
Vocational training	.2205	2.34**
Disability which limits work	.2805	2.02**
Married	-.0192	-.11
School age children	.1133	1.10
Prior migrant	.2036	2.08**
Occupation:		
Craftsman or kindred worker	-.0458	-.25
Operative (except transport)	-.1159	-.70
Transport equipment operative	.3118	1.57
Labor Market Characteristics:[b]		
Employment growth, 1965-1970 (%)	.1165	2.07**
Unemployment rate (%)	-.0083	-.13
Population (10^6)	.0219	1.46
Population density (10^3)	.3191	2.30**
Industry Characteristics:[c]		
Employment growth, 1965-1970 (%)	.0255	1.30
Percent union	.0058	.57
Weekly earnings[d]	-.0309	-1.80*
Fatal injury risk	2.8831	2.87***

* t-test significant at the 0.10 level.
** t-test significant at the 0.05 level.
*** t-test significant at the 0.01 level.

[a]All variables are defined in the text. Estimates were obtained using Group A microdata (4,511 observations) also described in the text. Industry switching was based upon three-digit Standard Industrial Classification codes for industries of employment in 1965 and 1970. The log likelihood ratio test statistic was significant at the 1 percent level.

[b]Based upon state of residence in 1965. Variables represent 1965 values except where noted.

[c]Based upon industry of employment in 1965 and national data. Percent union and fatal injury risk represent 1968 and 1969 values, respectively.
[d]Represents W* in equation (2). Predicted 1965 values were obtained from an estimate of equation (1) employing Group A microdata. See the text.

attenuates with age). In addition, industry switching is also responsive to labor market conditions, such industrial mobility being increased by rapid local job creation (employment growth) as well as compactness of local employment opportunity (population density), and thus lower search (transactions) costs. Finally, interindustry mobility among white-male manufacturing workers is augmented and diminished by risk of fatal injury in the workplace and by weekly earnings received, respectively.[12] Thus, to the extent that additional earnings provide insufficient compensation for added risk (based upon workers' willingness to pay for risk deduction), individuals likely "vote with their feet" to find employment in other, less risky, industries.

4.3 The Wage-Risk Trade-Off: Market Price vs. Willingness to Pay

As stated above, there are reasons to suspect that the market price of risk reduction [W'(R)] diverges from workers' willingness to pay for this reduction (dW/dR). Thus, safety in the workplace may quite possibly be provided in suboptimal amounts. Based upon the econometric estimates above, we are now equipped to investigate this question of market failure empirically.

Notice that W'(R) can be derived from coefficient estimates in Table 3 on the fatal injury risk variable and its interaction with percent union, while dW/dR can be determined by equation (3) from estimates on weekly earnings and risk in Table 4. Also note that an estimate of the total risk reward (compensatory earnings) provided workers at the prevailing level of risk exposure (say R_2) is equal to $W'(R_2) \cdot R_2$. On the other hand, workers' willingness to pay, through foregone earnings, for the elimination of all industrial work hazards (such that $R = 0$) is equal to $(dW/dR) \cdot R_2$, the derivative being evaluated at R_2.[13] Termed collectively "fatal injury risk rewards," these terms were

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[12] Weekly earnings in Table 4, W* in equation (2), represent the predicted 1965 value for each manufacturing worker obtained from an estimate of equation (1) employing Group A microdata. See the section on wage (earnings) determination below.

[13] Since a "one unit" change in fatal injury risk (a fraction) is difficult to interpret, comparisons below of market price and willingness to pay for risk reduction, W'(R) and dW/dR respectively, will be made on the basis of these wage-risk trade-offs

evaluated on the basis of estimated parameters in four earnings equations (two of which are shown in Table 1) as well as estimates from the industry switching equation. Results are listed in the first column of Table 5.

Based upon estimates of the earnings equations for white-males employed within manufacturing, weekly earnings premiums for risk, evaluated at "market" prices, range from $1.91 to $3.28 (in 1965 dollars). Such "compensating differentials" comprise 1.7 -2.9 percent of weekly earnings (see column 2). Alternatively, the estimates in the first column in Table 5 also represent the weekly wage consession <u>required</u> of manufacturing workers to reduce risk to zero. However, an estimate obtained from the industry switching equation with equivalent data indicates that such workers are <u>willing</u> to forgo $4.64 per week (4.0 percent of their income) to eliminate risk of fatal injury from the workplace. Thus, willingness to pay for risk reduction within manufacturing exceeds its market price (by a minimum of 41 and a maximum of 143 percent).

Since willingness to pay for risk reduction exceeds the market price of providing this reduction, implied value of life estimates derived from wage/earnings (or hedonic) equations are downward biased. Examples of such estimates derived from the four wage equations are provided in the last column of Table 5.[14] In addition, the fifth entry in column 3 represents implied value of life based upon willingness to pay for risk reduction (as determined from the industry switching equation). This and the other four value of life estimates were determined in accordance with accepted practice. Assuming for an average worker an average work week of 40 hours and an average of 50 weeks worked per year (two thousand hours), the fatal injury risk variable (fatalities per million hours worked) represents the likelihood of an adverse outcome per 500 workers. Thus, multiplication of $W'(R)$ or dW/dR by 500 yields the value that workers place on their life for small changes in the likelihood of death on-the-job. In addition, this value must be inflated by a factor of 50 to convert weekly earnings [employed in both equations (1) and

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weighted by the mean risk level. See the footnotes in Table 5. Theoretically, that such an approach may overvalue (undervalue) estimates of risk compensation based upon willingness to pay (market price) has been noted by Smith (1979).

[14] These estimates are consistent in magnitude with other implied value of life estimates obtained from wage/earnings equations and an equivalent job hazard variable. In this respect, see Viscusi (1978).

TABLE 5

Fatal Injury Risk Rewards in Manufacturing and Implied Value of Life (1965 Dollars).

Source of Estimate:	Fatal Injury Risk Reward: Weekly Premium[c]	Relative Weekly Premium[d]	Implied Value of Life[e] ($\times 10^6$)
Earnings Equation/ by dependent variable:			
1969 Earnings[a]	$1.91	1.7%	$ .972
Ln (1969 Earnings)[a]	2.45	2.1	1.250
1965 Earnings	2.28	2.0	1.145
Ln (1965 Earnings)	3.28	2.9	1.644
Industry Switching Equation[b]	4.64	4.0	2.330

[a] See equation (1) and Table 3. 1969 earnings were deflated to 1965 dollars based upon average weekly earnings of production workers on manufacturing payrolls (deflator of .8303). For a similar procedure, see Dillingham (1985). For example, $2.45 is the product of .8303, $60.20, and the mean risk level .049. The value of $60.20 is obtained from the mean of weekly wages ($173.63), the mean union rate, and the estimated equation as mean wage times the sum of the coefficient on fatal injury risk (1.6521) and the coefficient on the interactive variable (-.0293), the latter multiplied by the mean union rate (44.55).

[b] See equations (2) and (3), and Table 4. Evaluated at the mean risk level. 1965 dollars.

[c] This premium is equal to W'(R) in equation (1), or dW/dR in equation (3), multiplied by the mean fatal injury risk level.

[d] This is calculated as 100 (weekly premium)/mean 1965 weekly earnings. ($115.04)

[e] The implied value of life is determined as W'(R) in equation (1), or dW/dR in equation (3), multiplied by 25,000 (which represents the conversion of the risk data per 500 workers times 50 weeks per year). For example, .972 = 25000 x (175.9491 - 2.8983 x 44.55) x .8303.

(2)] to annual compensation.[15] Based upon the mean of the first four entries ($1.253 million dollars), implied value of life estimates obtained from wage/earnings equations and, thus, market evaluations of risk rewards, understate life's value by 46 percent. Compared to the government agencies listed in Table 1, the implied value of life from

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[15] Important income effects overlooked by this procedure are examined by Viscusi (1978b).

the industry switching equation (when inflated to current dollars) tends to agree only with the higher values utilized by OSHA and the EPA.

5 CONCLUSIONS

Public policy debates over environmental regulations for health and safety hazards have often centered on cost-benefit analysis in which the critical parameter is the implied value of life. This necessity to evaluate environmental policy affecting human life in dollar terms is becoming increasingly common in both the United States and Europe. For example, in the United States the FAA proposal in 1985 for all airline manufacturers to strengthen seats depended critically on the assumed value of life estimates. In this paper we have examined the valuation of life from the economist's perspective of compensating wage differentials. Empirical studies of compensating wage differentials attributable to risk in the workplace assume, most often implicitly, that workers' willingness to pay for risk reduction (safety) is equal to the market "price" of providing this reduction. Thus, it is assumed that labor markets are observed in a state of pure competition in long-run equilibrium. In this regard, workers and their employers are believed to possess perfect information regarding work hazards, the likelihood of adverse outcomes stemming from these hazards, and the cost of providing additional safety in the work place.

However, to the extent that such information is not known in its entirety, the prices at which risk is "bought" and "sold" within industrial labor markets are likely to differ. Thus, evaluations of the wage-risk trade-off will vary to the extent that the market (employer's) price diverges from workers' willingness to pay. Under the assumption that manufacturing workers "vote with their feet" among jobs in various industries in response to working conditions and compensation received (and thereby reveal their workplace preferences in the process), this willingness to pay for risk reduction was shown to exceed, by 41 to 143 percent, the market price of incremental safety.

Since value of life estimates derived from wage/earnings (hedonic) equations also assume equivalence between the willingness to pay for, and market price of, added safety in the workplace, such valuations are downward biased. A mean estimate of this bias was placed at 46 percent. Finally, by employing workers' own wage-risk trade-offs at prevailing levels of risk exposure within manufacturing, the implied value of life for white-male workers was shown to exceed two million (1965) dollars. These results suggest that the use of "traditional" value of life estimates in formulating environmental regulations may be too low.

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