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Compensating Wage Differentials for Fatal Injury Risk in Australia, Japan, and the United States

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Abstract

Our research infers the effects of institutionalized wage setting and lengthy worker–firm attachment by comparing estimated compensating wage differentials for fatal injury risk in Japanese, Australian, and U.S. manufacturing. Hedonic labor market equilibrium regressions for Japan reveal a statistically fragile compensating wage differential of 0% to 1.4% for exposure to the average fatality risk compared to employment in a perfectly safe workplace. Australian workers receive a statistically robust 2.5% estimated wage premium. Using new data on work-related fatalities, we find a 1% compensating wage differential in U.S. manufacturing that becomes more positive and statistically less significant as data are aggregated.

Policymakers are reluctant to use estimated compensating wage differentials for health risks in designing programs to reduce environmental hazards or encourage workplace safety because the estimates vary widely (Viscusi, 1983, Chapter 6; Moore and Viscusi, 1988). Researchers have tried to understand the divergence across studies of compensating wage differentials and the implied value of a worker's life by focusing on parameter robustness to changes in functional form and risk measures using data for a single country's labor markets (Dillingham, 1985; Marin and Psacharopoulos, 1982; Leigh and Folsom, 1984; Moore and Viscusi, 1988; Olson, 1981). We take another tack and examine three countries' labor markets, each one having a feature likely to influence compensating wage differentials. In particular, by comparing estimated hedonic labor

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market equilibrium equations for Australia, Japan, and United States, we clarify the effects of institutionalized wage setting, lengthy worker-firm attachment, and data aggregation on compensating wage differentials for fatal-injury risk.

Union workers in the United States earn a larger premium for exposure to workplace hazards than similar nonunion workerss (Thaler and Rosen, 1975). The effect of unionization on the compensating wage differential may reflect unions' concern with workplace safety or a larger proportion of union workers in high-risk industries. If unions negotiate contracts with workplace hazards in mind, then a highly unionized country, such as Australia, would have a higher estimated wage premium for injury risk than the United States. Alternatively, if union workers are just overrepresented in high-risk industries in the United States, then Australia, with its more equal unionization across industries, would have a lower wage premium for injury risk.

U.S. workers are rewarded for the type of job performed, and interfirm mobility and occupational identification contribute to a positive relationship between wages and the likelihood of a work-related injury or disease. The Japanese economy rewards workers for long-term attachment, since larger employers use internal labor markets and life-cycle measures, including age or years of service, to allocate training and promotion opportunities. Workers in the larger Japanese firms are also less tied to a particular occupation and more tied to a specific firm than workers in the United States, so a wage premium for workplace health hazards may not exist in Japan.

The extensive union and government involvement in wage setting in Australia and the relatively low interfirm worker mobility in Japan make us expect a higher compensating wage differential in Australia and a smaller compensating wage differential for injury risk in Japan than in the United States. Single-equation estimates of the hedonic equilibrium wage locus for two-digit manufacturing in Australia, Japan, and the United States yield the anticipated ranking of compensating wage differences.

The larger Japanese manufacturing firms pay higher wages and have safer workplaces; there is little econometric evidence of a positive relationship between wages and the rate of fatal injuries across two-digit manufacturing industries in Japan. The aggregate manufacturing data for Australia permit a more extensive list of independent variables including the ability to control for the effects of interstate differences in workers' compensation generosity and interindustry differences in firm size. We find that Australian manufacturing workers exposed to the mean fatality risk earn 2.5% higher wages than Australian workers will earn in a completely safe industrial setting. The 2.5% estimated compensating wage differential in Australia is robust to changes in the list of independent variables and estimation technique.

The U.S. Current Population Survey data are richer than the aggregate Australian and Japanese data in a number of ways—more independent variables and a larger sample size with potential for disaggregation—and can be matched to new measures of job-related injury risk produced by the National Institute for Occupational Safety and Health. We find a statistically significant 1% compensating wage differential for exposure to mean fatality risk in the individual manufacturing data. The estimated compensating wage differential in manufacturing rises, but loses statistical significance, with aggregation to states, regions, and the United States overall. A final benefit of the U.S.

data is the spacial variability of potential workers' compensation insurance payments. Our econometric results imply that the insurance for injuries implicit in compensating wage differentials trades off against formal insurance in the form of workers' compensation. We estimate that the U.S. workers' compensation system as it now exists lowers the compensation for exposure to the mean fatality risk in manufacturing from 11% to 1%.

1. Theoretical background

Our empirical research rests on the theory of compensating wage differentials in long-run labor market equilibrium. The eventual sorting of workers and firms in the labor market creates an equilibrium locus of joint wage–workplace safety outcomes conditioned by the characteristics of suppliers of labor, the characteristics of the demanders of labor, and elements of the institutional and legal environment. Although workers may not have enough information when initially accepting employment to establish compensating differentials for work-related health hazards, they eventually learn the true risks and create compensating differentials in a competitive environment through interfirm mobility (Viscusi, 1979). Worker mobility is not necessary for compensating wage differentials to exist, though, because the government or a union can formalize higher pay for workers with more hazardous jobs.

By reducing the monetary loss from a workplace injury, workers' compensation insurance lowers the pay premium necessary for workers to accept workplace health hazards. Interstate variation in available benefits creates additional variation in compensating wage differentials. The Japanese workers' compensation insurance system is a national program with uniform benefits across prefectures (Williams, 1985, 1988). In the cases of Australia and the United States, we emulate researchers who have accounted for differences in the workers' compensation system across states by including a potential generosity measure when estimating compensation for exposure to work-related health hazards (Viscusi and Moore, 1987; Moore and Viscusi, 1988).

For some issues, researchers must uncover the structural equations underlying hedonic equilibrium—employer cost curves and worker indifference curves. Recent research establishes the stringent identifying restrictions needed to estimate the supporting indifference and cost curves in hedonic equilibrium models (Brown and Rosen, 1982; Epple, 1987). Alternatively, a researcher can simulate the complete model over a set of structural equation parameter values, including parameters representing public policy influencing job safety (Kniesner and Leeth, 1988, 1989a, 1989b). The issues we address need only econometric estimates of the hedonic locus, which is a reduced-form equation estimable with single-equation methods.

Hedonic labor market equilibrium is described algebraically by

$$\ln W = f(I \mid S, D, \mathrm{WC}, \epsilon), \tag{1}$$

where W is the wage rate and I is a vector of information concerning the likelihood of a work-related injury. The variables conditioning the equilibrium level of wages and the compensating wage differential, $\partial \ln W / \partial I$, include characteristics of the sellers of labor (S), characteristics of the buyers of labor (D), and characteristics of the workers' compensation system (WC). Note that we have included a stochastic error term (ϵ) to emphasize that the hedonic equilibrium relationship is inexact; we have also written the hedonic equilibrium in semilogarithmic form, the most popular specification in econometric research.

Estimates of the hedonic labor market equilibrium locus for the United States using the Bureau of Labor Statistics' fatality rate measures and micro-cross-section data on wages and worker–firm characteristics show a 2% to 4% annual wage premium on jobs with the mean risk of a fatal work-related injury compared to jobs with zero risk (Smith, 1979; Viscusi 1983, Chapter 6). The estimated compensating differential is reduced by potential workers' compensation payments (Viscusi and Moore, 1987). Our research uses manufacturing data to estimate compensating wage differentials and how they are affected by potential workers' compensation payments. We selected Australia and Japan to compare to the United States because they have labor market features that interest U.S. economists and policymakers.

The Japanese worker in a large firm gets more on-the-job training than the typical U.S. worker, which contributes to Japan's higher rate of economic growth (Mincer and Higuchi, 1988). A feature of the Japanese economy receiving less attention is the comparatively low worker interfirm mobility supporting the comparatively high on-the-job training accumulation. Postwar monthly worker separation rates are two to three times higher in U.S. manufacturing than in Japanese manufacturing (Mincer and Higuchi, 1988). Because worker mobility helps establish compensating differentials for workplace risks, we expect the estimated value of $\partial \ln W / \partial I$ to be smaller in Japan than in the United States.

The Australian labor market interests economist and policymakers because of its highly institutionalized wage-setting process involving both unions and government. There is a channel in Australia for establishing formal salary supplements, known as dirt pay, for exposure to unpleasant or dangerous working conditions (Brooks, 1988; Jones, 1988). Evidence from the United States indicates a larger premium for workplace hazards in the union sector, which can reflect unions' concerns with workplace health and safety issues or overrepresentation of union workers in riskier U.S. industries. The estimated hedonic wage locus for Australia, with its more uniform unionization and highly institutionalized wage setting across industries, should clarify the effect of unions on compensating wage differentials. Specifically, comparing the relative steepness of the estimated hedonic loci in Australia and the United States may help us distinguish between competing interpretations of the effect of unions on compensating wage differentials for injury risk.

The final step of our econometric work will be to estimate the hedonic locus for U.S. manufacturing. We produce our own estimates because we want to compute compensating wage differentials with a new fatality rate measure and to identify the effect of aggregation on compensating wage differences. Our Australian and Japanese data are aggregate two-digit manufacturing cross sections, whereas our U.S. data are a cross section of individuals from the Current Population Survey. Comparing the regression results using individual U.S. data with results from the U.S. data aggregated to mimic the Australian and Japanese data should clarify the econometric effects of aggregation and suggest how estimated compensating wage differentials might change in Japan and Australia were they also estimated with individual data.

2. Empirical results—Japan

Because production levels affect injury rates and the equilibrium tradeoff between wages and injury risk, we have selected years for Australia, Japan, and the United States that are neither the troughs nor the peaks of business cycles. Regressions for Japan use public-use cross-section data, which cover firms with at least 30 employees in the 21 two-digit manufacturing industries in 1986 (*Yearbook of Labour Statistics, 1986*). Japan was experiencing a slight economic slowdown in 1986. Although the unemployment rate was unchanged from the year before, real GNP growth was lower, 2.4% in 1986 versus 4.9% in 1985. Inflation fell to 0.6% in 1986 from 2.0% in 1985, and the annual growth rate of nominal earnings in manufacturing fell to 2% in 1986 from 4% in 1985 (*Yearbook of Labour Statistics, 1986*).

Summary statistics for all regression variables appear in table 1. The dependent variable in all Japanese regressions is the logarithm of average monthly earnings, which averages about 300,000 yen or $1800.[1] The few independent variables available to represent the underlying conditioning characteristics of workers and firms in the two-digit industry public-use data for Japan are sex composition of the labor force, hiring and separation rates, and industry size. The *Yearbook of Labour Statistics* is richer in measures of workplace health hazards; it contains the rates of injuries overall, fatalities, permanent total disabilities, and temporary total disabilities. Work-related injuries are less frequent in Japan than in the United States. Based on Bureau of Labor Statistics data, the 1986 fatality rate in the United States is 1 per 10,000 manufacturing workers versus 0.3 per 10,000 manufacturing workers in Japan.

Table 2 reports weighted least squares estimates using the inverse of industry employment to adjust for heteroskedasticity.[2] The first two lines are bivarate regressions between injury rates and wages, which establish a baseline. Although wages vary inversely with both the overall injury rate and the fatality rate, only the coefficient of the injury rate is statistically significant at the 0.1 level. When independent variables are added to control the effects of sex composition, new hires, and separations on wages, the coefficient of the overall injury rate changes little. When both the overall injury rate and the fatality rate appear in the regression (table 2, line 4), the coefficient of the fatality rate is positive and significant at the 10% level using a one-tail test. The fatality rate regression coefficient indicates that workers exposed to the average Japanese manufacturing fatality rate receive 0% to 1.4% higher pay than they would in a perfectly safe Japanese manufacturing industry.

We conclude that the compensating wage differential is at most 1.4% in the two-digit Japanese manufacturing data because the coefficient of the fatality rate is fragile. The fatality-rate coefficient becomes negative and insignificant if we alter the measures of the hiring and separation rates, delete the overall injury rate from the regression, or use

Table 1. Summary statistics

Country/Variable	Mean	Standard deviation	Max.	Min.
Japan 1986, N = 20				
Injury rate (per 100 workers)	0.411	0.275	1.232	0.084
Fatality rate (per 100 workers)	0.0032	0.0046	0.018	0.0
Proportion female production workers	0.230	0.177	0.709	0.0
New hire rate (per 100 per mo.)	1.625	1.057	5.243	0.434
Separation rate	1.764	1.036	5.387	0.667
Earnings (yen/mo)	306993	72973	460355	165864
Australia 1984-1985 N = 44				
Injury rate (per 100 workers)	8.13	5.28	21.7	2.7
Fatality rate (per 100 workers)	0.014	0.022	0.12	0.0
WC benefit (death, $A)	49433	4265	52993	42390
Employees per firm	50.4	42.7	204.5	17.1
Proportion female	0.24	0.16	0.73	0.06
Change in inventories ($A million)	82.1	117.1	559.3	−47.3
Materials purchases ($A million)	3773.6	4861.6	19955.9	448.2
Value added ($A million)	2114.9	2087.1	9569.6	483.6
Investment	197.3	390.6	2532.5	18.6
Earnings ($A/year)	18467	2963	24630	12278
U.S.A. 1978, N = 8868				
Injury rate (lost workday accidents per 100 workers)	5.46	2.32	11.1	2.2
Fatality rate (per 100 workers)	0.0436	0.044	0.621	0
WC benefit (death, max. per week, spouse)	168.27	67.51	654	88
Limit (on death benefit)	0.51	—	—	—
WCbenefit*limit	82.58	94.11	330	0
Nonwhite	0.12	—	—	—
Female	0.32	—	—	—
Union	0.36	—	—	—
New hire rate (avg. per mo. per 100)	3.05	1.08	5.3	1.4
Separation rate	3.85	1.52	7.7	1.7
Age	37.10	12.59	74.0	15.0
Education (years completed)	11.65	2.73	18.0	0.0
Married	0.75	—	—	—
Earnings ($/week)	256.27	133.74	999.0	70.0

Table 2. Compensating wage differentials in Japan and Australia

Country	Injury/Workers' comp. measures	Coefficient	P-value	Other independent variables	adj R^2
(1) Japan[a]	Injury rate	−0.520	0.0061	—	0.31
(2) Japan[a]	Fatality rate	−20.576	0.0556	—	0.14
(3) Japan[a]	Injury rate	−0.273[b]	0.0001	d	0.88
(4) Japan[a]	Injury rate	−0.325[b]	0.0001		
	Fatality rate	4.422[c]	0.1047	d	0.97
(5) Australia[e]	Injury rate	−0.0005	0.9112	—	−0.02
(6) Australia[e]	Fatality rate	2.1189	0.0611	—	0.06
(7) Australia[e]	Injury rate	−0.0058	0.0640		
	Fatality rate	1.8108	0.0290	f	0.65
(8) Australia[e]	Injury rate	0.0004	0.894		
	Fatality rate	1.729	0.0178		
	WC benefit	0.000014	0.0015	f	0.74
(9) Australia[e]	Injury rate	0.00063	0.8467		
	Fatality rate	10.6851	0.3373		
	WC benefit	0.000016	0.0031		
	Fatality rate * WC benefit	−0.00018	0.4194	f	0.73

[a]Dependent variable: log average monthly earnings in 1968, $N = 20$. Weighted least squares with (1/total employment) as weights.
[b]Changes little with alternative measures of separations and new hires or when estimated via ordinary least squares.
[c]Fatality rate coefficient goes to zero with changes in measures of separations and new hires or using ordinary least squares. Equals −8.199 with a P-value of 0.065 when injury rate is omitted and ordinary least squares used.
[d]Proportion of female production workers, new hire rate, and separation rate.
[e]Dependent variable: log annual earnings in 1984–1985, $N = 44$.
[f]Employees per firm, proportion women, change in inventories, materials purchased, value added, and investment.

ordinary least squares instead of weighted least squares as the estimation technique.[3] In addition, the coefficient of the overall injury rate is significantly negative in all regressions, which we interpret as indicating model misspecification. Specifically, the largest Japanese manufacturing firms have the lowest injury rates and pay the highest wages. Recent evidence from the United States emphasizes a firm-size effect on wages (Brown and Medoff, 1989). A defect in the regressions for Japan is our inability to hold firm size constant, which can create a negative omitted variable bias on the coefficients of the injury rates. We therefore caution against using public-use data on two-digit Japanese manufacturing industries in future research on wages.

3. Empirical results—Australia

In conjunction with the Australian Bureau of Statistics, each state publishes *Industrial Accidents* and *Manufacturing Establishments,* which provide the data for our Australian hedonic labor market equilibrium regressions. We study the 11 two-digit manufacturing industries in the four largest Australian states in 1984–1985. The states in our data—New South Wales, Victoria, Queensland, and South Australia—contain 90% of Australian manufacturing workers.[4]

Although the Australian economy was stagnating during the mid-1980s, there was an uptick during 1985. The unemployment rate declined to 7.9% from 8.3% in 1984, the inflation rate increased to 6% from 4.5% in 1984, and annual real GNP growth rate was the same in 1984 and 1985 (*Yearbook Australia 1988* and *Labour Statistics Australia 1986*).

Summary statistics of the variables in our Australian regressions also appear in table 1. The dependent variable is the logarithm of annual earnings, which averaged $A18,467 or $US14,774. Independent variables include average firm size, work-force sex mix, production levels, and nonlabor inputs. The set of conditioning variables is as complete as the aggregate Australian cross-section public-use data permit. Fatality rates in Australian manufacturing are closer to fatality rates in U.S. manufacturing than in Japanese manufacturing, averaging 1.4 per 10,000 workers. For comparison purposes, we discuss only regressions including the overall rate of injuries and diseases and the rate of fatal injuries and diseases, which are metered by Australian workers' compensation insurance claims.[5]

As in the United States, there is substantial interstate variation in available workers' compensation insurance benefits in Australia. To capture generosity differences simply and to use a generosity measure that is exogenous to individual worker behavior, we parameterize available workers' compensation benefits with each state's maximum benefits for a work-related death. The average maximum death benefit is $A49,400, or $US39,500, with the highest state benefit $A53,000 and the lowest state benefit $A43,000.

Baseline bivariate regressions for Australia appear in lines (5) and (6) of table 2. A similarity to Japan is that wages vary inversely with the overall injury rate, but the relationship is not significant statistically at conventional levels. In contrast to Japan, the fatality-rate coefficient is positive and significant ($p = .061$) in the baseline bivariate regressions for Australia.

One of the attractions of the Australian data is an ability to check whether workers' compensation benefits trade off against wages in hedonic equilibrium (to test whether the estimated value of $\partial \ln W/\partial \mathrm{WC}$ is negative) and to check whether the compensating wage differential falls with workers' compensation benefit generosity (to test whether the estimated value of $\partial^2 \ln W/\partial I \partial \mathrm{WC}$ is negative). In regression (8) of table 2, there is a positive partial relation between benefit generosity and wages, which we attribute to a simultaneous equations bias that we cannot remedy when using the aggregate cross-section data for Australia. Specifically, prosperous high-manufacturing-wage states also set up more generous workers' compensation systems.[6] Although statistically insignificant, the estimated interaction effect between the compensating wage differential for injury risk and potential workers' compensation benefits in Australia is negative in regression (9) of table 2.[7]

Finally, the regressions in table 2 say that manufacturing workers exposed to the mean Australian fatality risk receive a wage premium of about 2.5%. Unlike Japan, the 2.5% compensating differential in Australian manufacturing is robust to changing the list of control variables, including deleting the possibly endogenous measure of workers' compensation benefit generosity and replacing it with state dummy variables.

4. Empirical results—the United States

The Current Population Survey is the only micro cross section with enough individuals to examine the effects of data aggregation. We chose May 1978 because of its location in the business cycle and available ancillary data on worker union status and turnover. To elaborate, 1978 is at the middle of the 1975–1980 economic expansion. The unemployment rate was 6% in 1978 compared to 8.3% in 1975. Real GNP grew at a healthy 5.3% compared to −1.3% in the trough year. Inflation had also fallen to 7.6% in 1978 from 9.1% in 1975. Although later years have a similar position in the business cycle, only 1978 allows us both to identify union membership and to merge data on worker turnover rates (new hires, separations).

The Bureau of Labor Statistics no longer releases industry fatality rates at other than the one-digit level.[8] Our measures of workplace health hazards instead include average industry workdays lost to injuries and the fatality rates of the National Institute for Occupational Safety and Health, which are derived from a census of all occupational fatalities recorded on death certificates from 1980 to 1985.[9] Although the National Institute's fatality data are publicly available for only one-digit industries, they are disaggregated by state and provide an excellent picture of long-run interstate differences in fatality risk.[10] Moore and Viscusi (1988) prefer the National Institute's fatality data to those of the Bureau of Labor Statistics because the former are based on a census rather than a survey and are therefore freer of error. Moreover, the focus on interstate differences in the National Institute's data permits a more precise match of death risk with available workers' compensation benefits. Regression estimates of compensating wage differentials across one-digit industries using the National Institute's fatality data are larger than the typical regression estimates of compensating wage differentials using the Bureau of Labor Statistics' fatality data (Moore and Viscusi, 1988).

The dependent variable in our regressions with U.S. data is the logarithm of average weekly earnings of full-time manufacturing workers.[11] Independent varibles include race, sex, marital status, age, education, region, and the industry's average monthly new hire and separation rates. We again represent workers' compensation insurance generosity by a state's maximum weekly death benefits to the surviving spouse and include a dummy variable for whether the state had a maximum total spouse benefit.[12] We consider three levels of aggregation above the individual: states, regions, and two-digit manufacturing industries.[13]

The first block of regressions in table 3 illustrates the effect of aggregation on the estimated baseline compensating wage differentials. The fatality rate coefficient becomes more positive and less significant as the level of aggregation increases. The second block of four regressions in table 3 includes a set of control variables intended to mimic the Japanese regressions of table 2. Unlike Viscusi and Moore (1987), who find a weak tradeoff between potential workers' compensation insurance benefits and wage levels,

Table 3. Compensating wage differentials in the United States (see text for details)

Level of aggregation	Injury/workers' comp. measures	Coefficient	*P*-value	Other independent variables	adj R^2
Individuals[a]	Injury rate	0.0140	.0001	—	0.0047
Individuals	Fatality rate	−0.2748	.016	—	0.0005
Individuals	Injury rate	0.01412	.0001		
	Fatality rate	−0.2958	.0091	—	0.0054
States[b]	Injury rate	−0.00363	.503		
	Fatality rate	0.2572	.138	—	0.0009
Regions[c]	Injury rate	−0.00526	.665		
	Fatality rate	0.9340	.268	—	−0.0072
Nation[d]	Injury rate	0.00376	.862		
	Fatality rate	5.1219	.142	—	0.0214
Individuals	Injury rate	0.01905	.0001		
	Fatality rate	−0.03272	.735		
	WC benefit	0.0006	.0001		
	WC benefit * limit	0.0001	.0350	e	0.326
States	Injury rate	−0.00740	.153		
	Fatality rate	0.1812	.163		
	WC benefit	0.0009	.0001		
	WC benefit * limit	0.000008	.945	e	0.466
Regions	Injury rate	−0.01594	.114		
	Fatality rate	0.3387	.447		
	WC benefit	0.0014	.008		
	WC benefit * limit	0.00026	.506	e	0.789
Nation	Injury rate	−0.00107	.599		
	Fatality rate	0.0268	.232		
	WC benefit	0.0021	.466		
	WC benefit * limit	0.00055	.860	e	0.903
Individuals	Injury rate	0.0185	.0001		
	Fatality rate	0.1365	.1299		
	WC benefit	0.00037	.0001		
	WC benefit * limit	0.000014	.734	f	0.512
States	Injury rate	0.000331	.940		
	Fatality rate	0.0888	.456		
	WC benefit	0.00081	.0001		
	WC benefit * limit	0.000023	.827	f	0.623
Regions	Injury rate	0.00205	.784		
	Fatality rate	0.2920	.446		
	WC benefit	0.00019	.643		
	WC benefit * limit	0.00098	.042	f	0.906
Nation	Injury rate	−0.00772	.749		
	Fatality rate	0.5168	.850		
	WC benefit	0.00012	.645		
	WC benefit * limit	−0.0021	.647	f	0.961

Table 3. (continued)

Level of aggregation	Injury/workers' comp. measures	Coefficient	*P*-value	Other independent variables	adj R^2
Individuals	Injury rate	0.01782	.0001		
	Fatality rate	2.5141	.0001		
	WC benefit	0.00075	.0001		
	WC benefit * limit	0.00027	.0001		
	Fatality rate * WC benefit	−0.0121	.0001		
	Fatality rate * WC benefit * limit	−0.0056	.0001	f	0.514
States	Injury rate	0.0004231	.923		
	Fatality rate	3.4774	.0001		
	WC benefit	0.0013	.0001		
	WC benefit * limit	0.0003	.0254		
	Fatality rate * WC benefit	−0.0190	.0001		
	Fatality rate * WC benefit * limit	−0.0049	.0273	f	0.636
Regions	Injury rate	−0.000696	.923		
	Fatality rate	1.0490	.784		
	WC benefit	−0.00103	.342		
	WC benefit * limit	0.00281	.0006		
	Fatality rate * WC benefit	0.00034	.157		
	Fatality rate * WC benefit * limit	−0.00044	.005	f	0.915
Nation	Injury rate	0.01012	.631		
	Fatality rate	117.1689	.266		
	WC benefit	−0.0060	.674		
	WC benefit * limit	0.0778	.197		
	Fatality rate * WC benefit	0.1648	.601		
	Fatality rate * WC benefit * limit	−2.0370	.192	f	0.990

[a]Dependent variable: log average weekly earnings in 1978; $N = 8868$, ordinary least squares.
[b]$N = 682$, ordinary least squares.
[c]$N = 80$, ordinary least squares.
[d]Two-digit manufacturing industries, $N = 20$, ordinary least squares.
[e]Nonwhite, female, union, new-hire rate, separation rate.
[f]Age, age^2, nonwhite, female, union, education, region, marital status, new-hire rate, separation rate.

the coefficient of available insurance benefits is positive in all regressions in the second, parsimoniously specified, block of regressions in table 3.[14] Although the estimated fatality-rate coefficient again rises with the level of aggregation, the estimated compensating wage differential approaches statistical significance at conventional levels in only the state level regressions.

By comparing the second and third blocks of regressions in table 3, we see how the results change when an extensive set of independent variables replaces a list similar to the Japanese regressions of table 2. The effects of aggregation are now clarified. Because the averaging process removes noise from the data, adjusted R^2 rises espectedly with the level of aggregation. The coefficient of the fatality rate and its p-value also rise with aggregation. Reduced estimated coefficient precision with aggregation is the net result of two opposing forces: aggregation lowers both the regression's residual variance and the variances of the independent variables. In our case, the reduced variance of the independent variables dominates, lowering statistical significance with aggregation. Only the fatality-rate coefficient estimated with individual data is significant at the 0.1 level using a one-tail test, indicating a compensating wage differential for exposure to the mean fatality risk compared to employment in a completely safe manufacturing workplace of 1%.

Moore and Viscusi (1988) find a large reduction in the compensating wage differential with generosity of available workers' compensation benefits. The final group of regressions for the United States targets how potential insurance benefits affect the compensating wage differential for fatality risk in manufacturing. The coefficients of the regression using individual data indicate that if there were no workers' compensation insurance death benefits, then the estimated compensating wage difference in manufacturing would be 11%. At the means of the fatality rate and workers' compensation benefits-fatality interactions, the estimated compensating wage differential is 1%.[15]

Finally, the relatively small estimated compensating wage differential for death risk produced by the coefficients in table 3 is the result of our focus on manufacturing. By omitting other industries we exclude workers at the extremes of the risk spectrum. Specifically, in the Current Population Survey data the death rate in mining is 32 per 100,000 employees while the death rate in wholesale trade is 1 per 100,000 employees. When we include all workers in the eight nonagricultural one-digit industries in a hedonic wage regression, the estimated compensating wage differential for fatality risk rises to 3.4% (4.3% in the absence of available workers' compensation death benefits).[16] The 3% to 4% compensating wage differential across all industries is similar to other researchers' results using individual wage data and the National Institute's fatality rates (Moore and Viscusi, 1988).

5. Conclusion

Our research estimates the compensating wage differential for the risk of a work-related fatality in the manufacturing sectors of Australia, Japan, and the United States. The distinguishing features of the Australian and Japanese labor markets make us expect a lower wage differential in Japan and a higher wage differential in Australia than in the United States. Our econometric research produced three findings of note.

First, the relative ranking of estimated compensating wage differences is as expected. Because the estimated compensating wage difference is fragile in the regressions for Japan using aggregate two-digit manufacturing data, we conclude that the Japanese differential is not significantly different from zero. Using new data on work-related fatalities, we find a compensating wage differential in U.S. manufacturing of about 1%. Regressions for Australian manufacturing yield an estimated compensating wage differential of approximately 2.5%.

Second, we were able to aggregate workers in the U.S. manufacturing sector by state, region, and two-digit industry. The effects of aggregation are severe; estimated compensating wage differentials for fatality risk become more positive and statistically less significant as the level of aggregation increases.

Finally, researchers need to incorporate potential workers' compensation insurance benefits in hedonic labor market equilibrium regressions designed to estimate compensating wage differentials for workplace injury risks. We find that formal insurance for work-related fatalities substitutes for the insurance implicit in compensating wage differences in the U.S. manufacturing sector. Our estimates are that if there were no death benefits under workers' compensation insurance, the compensating wage differential would be 11% instead of 1%.

Notes

1. Separating earnings into contractual versus bonus components did not alter the empirical results.
2. A Goldfeld-Quandt test rejects the null hypothesis of homoskedasticity at the 10% level.
3. The fragility of the Japanese regression coefficients may also reflect small sample size. We therefore doubled the number of observations by combining data on the 20 manufacturing industries for 1984 and 1986. The results are similar to the coefficients reported in table 2. When we control for the proportion of women production workers, the new hire rate, and the worker separation rate in a weighted least squares regression, the fatality rate coefficient is significantly positive at the 10% level. However, the size of the estimated compensating wage differential for workers exposed to the mean fatality risk drops in half, to 0.73%. In a manner to our previously discussed results, the fatality rate coefficient is sensitive to the control variables used in the regression and to the use of weighted versus ordinary least squares. In no other specification, including ones incorporating time trends and industry (fixed) effects, is the fatality rate coefficient statistically significant; in many specification the fatality rate coefficient is negative.
4. The remaining 10% of manufacturing workers are in Tasmania, the Northern Territory, or in the Australian Capital Territory and are ignored because of missing data on work-related injuries.
5. See Kniesner and Leeth (1989a) for a numerical simulation of the difficulties with using workers' compensation claims data to study work-related injuries and compensating wage differentials.
6. Including a set of state indicators could potentially separate the effects of state-specific factors and workers' compensation benefits on wages. More than one year of data is needed because of the collinearity between state dummy variables and state workers' compensation benefits. We expanded our data set to include information on New South Wales, Queensland, and South Australia for 1986. (Data for Victoria are unavailable for 1986.) Additionally, because of limitations in the 1986 data, other conditioning variables are limited to industry work-force sex-mix and product turnover (essentially sales) per establishment. The coefficient estimates from pooled data for New South Wales, Queensland, and South Australia for 1984–1985 and 1985–1986 mirror the results of table 2 in both sign and significance when the state dummies (fixed effects) are excluded. When state dummy variables appear in the regression, the fatality rate coefficient remains positive and significant but the workers' compensation benefits coefficient turns negative and insignificant. Hence, the results of our ancillary regressions provide some evidence that workers' compensation benefits in the regressions of table 2 are, at least partly, reflecting unmeasured state-specific factors.
7. Workers' compensation death benefits should affect only the wages of workers exposed to potentially fatal workplace hazards, implying that only the interaction term between benefits and fatality rates need be included in regression (9) (Viscusi and Moore, 1987). When the workers' compensation variable is excluded, the fatality rate coefficient is insignificantly negative and the interaction term between benefits and fatality rates is insignificantly positive. Because workers' compensation benefits may be proxying for latent state-specific factors, we also ran regression (9) excluding workers' compensation generosity but including state dummy variables, and the results of table 2 were maintained.
8. The data from the Bureau of Labor Statistics are derived from a survey of establishments intended to measure injury risk, not fatality risk. The Bureau formerly released data at the two-digit and three-digit levels to researchers requesting it. After a study completed in 1985 by an outside consulting group, the

Bureau concluded that fatality rates derived from their establishment survey are misleading. The Bureau no longer releases two-digit or three-digit fatality data.

9. We thank Dr. Nancy A. Stout of the Division of Safety Research, Injury Surveillance Branch, Data Analysis Section, The National Institute for Occupational Safety and Health for providing us with the data, which are described in Moore and Viscusi (1988).
10. In states with less than six total deaths in the manufacturing sector, the National Institute did not calculate a fatality rate, and we coded it as zero.
11. Average weekly hours are 35 or more.
12. We also ran regressions where the spouse benefit was coded as zero if the worker was unmarried. This had no effect on our conclusion.
13. Fatality rates vary across national two-digit manufacturing industries because there is an unequal number of workers in each industry-state cell.
14. In a manner similar to Australia, the workers' compensation benefit variable may be capturing unmeasured state-specifc factors.
15. We also examined the impact of workers' compensation benefits on the compensating wage differential for fatal injury risk in regressions that exclude the level of benefits as a separate variable. (see note 7). At the individual level, and at every level of aggregation, the fatality rate coefficient is insignificantly negative and the interaction term between the fatality rate and available workers' compensation benefits is insignificantly positive. We interpret the lack of statistical significance as reflecting unmeasured state-specific factors that are indirectly controlled by including both the level of benefits and the interaction of workers' compensation benefits and fatality rates in the regression. Because, in the case of the United States, we use only data for 1978 for reasons discussed at the beginning of section 4, and because our fatal-injury risk measure varies only by state, we cannot try to separate workers' compensation insurance effects from state-specific (fixed) effects, as we did for Australia, by including a set of state indicator variables as regressors.
16. We include all independent variables listed in the last set of regressions of table 3 except overall injury rate, new hire rate, and separation rate. The worker turnover rates are unavailable for non-manufacturing industries, and omitting the overall injury rate does not affect the estimated fatality rate coefficient. For comparison purposes we reestimated the hedonic wage function with individual data for manufacturing workers, excluding the overall injury and worker turnover rates; the estimated compensating wage differential for fatality risk increased to 1.6% in manufacturing.

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