Estimating the Value of Safety with Labor Market Data: Are the Results Trustworthy?

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Summary

We use a panel dataset of UK workers to look for evidence of compensating wage differentials for workplace risk. Risk data are available at the four-digit industry level or at the three-digit occupation level. We discuss various econometric problems associated with the hedonic wage approach, namely measurement error, instability of the estimates to specification changes, and endogeneity. We find that if we assume a classical measurement error, the true risk signal would be completely drowned out in our data, which would imply a severe downward bias of the OLS coefficient on risk. But this prediction is at odds with our OLS estimates of the VSL, which are large, especially for blue collar workers. Further, the coefficient on risk changes varies dramatically with the inclusion or exclusion of industry and/or occupation dummies, as well as with the addition of nonfatal risk. When we instrument for risk, which we treat as endogenous with wage, and apply 2SLS or a procedure suggested by Garen (1988), we find negative associations between risk and wages for all workers, which is against the notion of compensating wage differentials, or, for blue-collar workers, extremely large VSL figures. Finally, we exploit the panel nature of our data to apply various estimation procedures (the "within" estimator, GLS and the Hausman-Taylor procedure) that correct for unobserved heterogeneity and endogeneity. The coefficient on risk is usually negative and insignificant for the sample of all workers, which once again questions the notion of compensating wage differentials. For blue-collar workers we obtain reasonable VSLs, but the association between risk and wages is not statistically significant. We conclude that if compensating differentials for risk exist, measurement error, other econometric problems, and the changing nature of labor markets prevent us from observing them. We also conclude that models and techniques for panel data that account for unobserved heterogeneity and endogeneity seem more reliable than the techniques typically employed with cross-sectional data.

Keywords: Value of Life, Labor Market, Wage Hedonics

JEL Classification: D16, H43, J17, J28, J31

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I. Introduction

The Value of a Statistical Life (VSL) is a key input into the calculation of the mortality benefits of environmental policies or other safety regulations that save lives. The VSL is a summary measure of the rate at which individuals are prepared to trade off income for reductions in the risk of death. It can be equivalently described as the total willingness to pay by a group of N individuals experiencing a uniform reduction of 1/N in their risk of dying. The concept of VSL is generally deemed as the appropriate construct for ex ante policy analyses, when the identities of the people whose lives are saved by the policy are not known.

Ideally, to estimate the benefits of a policy that reduces the risk of dying, we should ask the beneficiaries of the policy how much they are willing to pay for the risk reduction. When this is not feasible, estimates of the VSL from other contexts are applied to the policy under consideration. Clearly, doing so assumes that individuals would apply the same marginal rate of substitution between income and risk in both the original and the new policy context.

Compensating wage studies are a commonly used approach to estimating the VSL. This approach uses data from labor markets, but the resulting estimates of the VSL are frequently transferred to the environmental policy context. For example, the US Environmental Protection Agency relied on 21 compensating wage studies, out of a total of 26 studies, to produce the VSL it uses in its policy analyses ($6.1 million 2000 dollars).[1]

The rationale of compensating wage studies is that workers must be offered higher wages for them to accept jobs with a greater risk of dying, *ceteris paribus*, and that employers are willing to do so to the extent that it is cheaper than installing safety equipment in the workplace. The demands of workers and the offerings of firms will meet at the tangency points between the workers' indifference curves and the firms' isoprofit curves in the risk-wage space. The hedonic wage relationship is the locus of these tangency points, and the slope of this locus is the willingness to pay for a marginal decrease in risk.

Viscusi and Aldy (2003) survey compensating wage studies and VSL estimates from all over the world. Based on US studies, they recommend VSL figures of $5-9 million. Estimates of the VSL based on compensating wage studies are available for several European countries, including the UK, where they usually range between $4 and $11 million, and Switzerland (6.5-9.5 million 2000 dollars; Baranzini and Ferro Luzzi, 2001).

Some recent work, however, has cast doubt on how credible these estimates are. Leigh (1995), Arabsheibani and Marin (2000, 2001), Black and Kniesner (2003), and Black et al. (2003) suggest that the estimates of the VSL from compensating wage studies are econometrically very fragile, for reasons that include poorly measured workplace

[1] This reliance on labor market estimates of the VSL is not uncontroversial. The use of VSL figures from compensating wage studies when computing the mortality benefits of environmental policies has been criticized on the grounds of the fact that it mirrors the preferences of healthy males whose average age is 40, rather than those of the primary beneficiaries of environmental policies—the elderly and those in compromised health. Adjustments to the VSL for remaining life years were subsequently proposed, and eventually repealed.

risk, collinearity of risk estimates with industry dummies used to account for inter-industry wage differentials, endogeneity (the level of risk facing an individual may be determined simultaneously with their job, and hence with the wage), omitted regressors, and heterogeneous preferences for risk and income.

We illustrate the severity of these problems and their effects on the VSL using a panel dataset of UK workers. To our knowledge, this is the first time panel data are used to study the robustness (or lack thereof) of VSL estimates in the labor market context. We conclude that our risk data is affected by measurement error, despite using industry risk data at the four-digit level (or occupational risk at the three-digit level). If the measurement error is assumed to be classical (i.e., uncorrelated with the regressors in the compensating wage equation), its variance is very large relative to the variance of the true risk variable, in which case the estimates of the compensating wage differentials would be affected by a severe downward bias. Most likely, however, as in Black and Kniesner (2003), the measurement error is non-classical, which means that the direction and magnitude of the bias are difficult to assess.

The panel nature of our data allows us to control for unobserved heterogeneity, even when the unobserved effects are correlated with included regressors (as in the "within" estimator, which is well suited to fixed effects model), and to develop alternate instrumental-variable techniques (e.g., Hausman and Taylor, 1981) to address endogeneity of risk with wages. When we use the sample of all male workers, we find negative, and usually insignificant, associations between risk and wages, which is against the very notion of compensating wage differentials.

When we restrict attention to male blue collar workers and use estimation techniques based on panel data the association between risk and wages is insignificant but the corresponding estimate of the VSL is within the range generally considered acceptable (Viscusi and Aldy, 2003). By contrast, estimates based on approaches that treat the data as independent cross sections produce disproportionately large VSL figures, even when they correct explicitly for endogeneity. Taken together, these results raise serious doubts about the credibility of VSL figures based on labor market studies—which are generally based on cross-sections of data—and underscore the importance of interpreting and using existing VSL estimates with caution in policy analyses.

The remainder of the paper is organized as follows. In section II, we present the concept of VSL and illustrate how it is usually estimated using labor market data. In section III, we discuss the main econometric limitations of the conventional approach. Section IV illustrates these limitations and explores possible remedies using UK worker data. Section V concludes.

II. VSL Estimates from Compensating Wage Studies

A. The Value of a Statistical Life

The VSL is the rate at which individuals are prepared to trade off income for risk reductions:

$$VSL = \frac{\partial WTP}{\partial R} \tag{1}$$

where R is the risk of dying and WTP is the Willingness to Pay for a reduction in risk, i.e., the maximum amount that can be subtracted from an individual's income to keep his or her expected utility unchanged for specified levels of risk.

The VSL is derived within an expected-utility maximization context, where individuals are assumed to derive utility from the consumption of goods. Formally, let U(y) denote the utility derived from income y when the individual is alive, V(y) the utility of income to the individual after he is dead,[2] and let R denote the risk of dying in the current period. Expected utility is expressed as EU=(1-R)·U(y)+R·V(y). Based on these assumptions, it can be shown that the VSL is equal to

(2) $$VSL = \frac{U(y) - V(y)}{(1-R) \cdot U'(y) + R \cdot V'(y)}$$

and that the VSL is positive as long as the utility of income is higher when one is alive.

B. Estimates of the VSL from Compensating Wage Studies

Figure 1 shows the hedonic wage curve, which is the locus of tangency between workers' indifference curves (EU) and the firms' isoprofit curves (OC) in the risk wage space. In a typical compensating wage study, data are gathered on the wage rate, education, experience, occupation, and other individual characteristics of workers and workplace characteristics. These data are then used to run a regression relating the wage rate to the risk of fatal and non-fatal injuries, while controlling for education and experience of the worker, and other job and worker characteristics thought to influence wages. A frequently used specification of the wage regression is:

(3) $$w_i = \beta_0 + \mathbf{x}_i\beta_1 + p_i\beta_2 + q_i\beta_3 + \varepsilon_i$$

where w_i is the wage rate for worker *i* and, x is a vector of individual, workplace and occupational characteristics, such as experience, education, age, gender, marital status,

[2] This utility would accrue from any bequest the individual plans to make. For simplicity, it is often assumed to be zero.

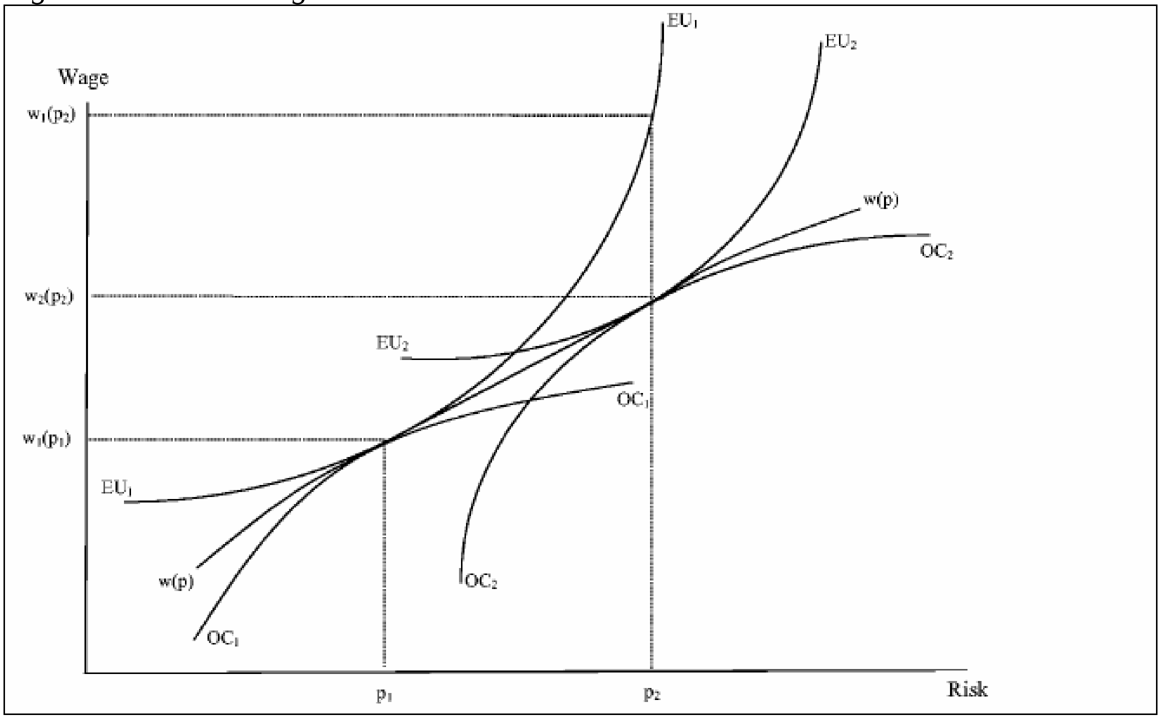
union status of the worker, industry dummies, occupation dummies, and location dummies. [3] The variable p measures the risk of dying on the job, while q is the risk of non-fatal injuries. The βs are the coefficients to be estimated, and the VSL can be inferred from β_1. For example, if w measures annual earnings and p annual fatal workplace risk in X per 100,000, then the VSL is calculated as $(100{,}000 \times \beta_2)$.

Viscusi (1993) argues that q must be included in the compensating wage equation. Since p and q are generally highly correlated, failure to do so would result in biased estimates of the β_1 coefficient, and hence of the VSL. Viscusi and Aldy (2003) examine some 60 wage-risk studies conducted all over the world, finding that many of them, however, do not control for non-fatal risk.[4]

What are the VSL figures typically estimated in compensating wage studies? As noted earlier, Viscusi and Aldy (2003) survey studies from all over the world, and for the US and recommend a range of \$5 to \$9 million (2000 US dollars). Recent compensating wage studies based on European labor markets (Siebert and Wei (1994); Sandy and Elliott (1996); Arabsheibani and Marin (2000); Sandy (2001)) peg the VSL in the range between €4.3 million and €74.4 million (equal to \$4.0 million to \$68.5 million at the 2000 exchange rate). A meta-analysis by CSERGE (1999) generates a range of VSL figures between €2.9 million and €100 million, resulting in weighted average equal to €6.5 million (all 2000 €, the corresponding dollar amounts being \$2.7, \$92.1 and \$6.0 million.)

[3] In empirical work, the logarithmic transformation of the wage rate often replaces *w* as the dependent variable in the regression. The wage rate, *w*, and fatality risk, *p*, are usually measured on an annual basis.
[4] In addition, Viscusi (1993) recommends that equation (3) should include expected worker compensation in the event of a non-fatal workplace accident, i.e., (WC×q), where WC is the level of workers compensation paid out to the worker if he experience an accident at work.

Figure 1. Hedonic wage curve



Source: Viscusi and Aldy (2003).

Using the 1995 Swiss Labor Force Survey (SLFS) and the 1994 Swiss Wage Structure Survey (SWSS), Baranzini and Ferro Luzzi peg the VSL implicit in the choices of Swiss workers in the range of 10 to 15 million Swiss Francs (6.5 to 9.5 million 2000 US dollars). They find that the VSL depends on risk level, union coverage, and age. Using Eurostat survey data for 1995, Barone and Nese (2002) find no evidence of a significant relationship between wage rates and objectively measured workplace risk in Italy.

III. Limitations of Compensating Wage Studies

Based on Leigh (1995), Black et al. (2002), and Black and Kniesner (2003), one suspects that compensating wage studies are rife with econometric problems. In this section, we discuss three main problems: (1) measurement error in the risk variable, (2) lack of robustness with respect to specification changes, and (3) endogeneity of risk or unobserved heterogeneity in wages and in the preferences for risk and income.

A. Measurement Error

The first key issue is how to measure p and q in equation (3). Job choice and wages should be driven by perceived as opposed to actual risk levels, yet researchers have typically measured p and q using official labor statistics (i.e., fatality rates within a certain occupation or industry).[5] Even if the objective risk were correctly measured (a matter we discuss below), merely allowing it to be the unbiased guess of workers' subjective risk estimate is sufficient to make p and q in equation (3) affected by (classical) measurement error. This would result in a downward biased estimate of β_2, and would understate the VSL.

Measuring objective risk, however, is far from simple. Compensating wage studies typically assign to worker i the fatality or non-fatal accident rate of his or her industry or occupation. Doing so ascribes the same level of risk to a miner and a secretary in a mining company (if the industry-level risk is assigned to all workers in that industry), or to factory workers in all industries and establishments (if occupation risk rates are assigned to all workers with that occupation). Both assumptions are unrealistic, unlikely to mirror actual individual worker risk, and create a problem of errors in variables. Formally, instead of observing p_i, we observe

[5] We are only aware of two studies that attempted to regress wages on perceived, rather then objective, risk (Gegax et al., 1991, and Lanoie et al., 1995). The Gegax et al. study, however, asked respondents to identify their own risk of dying on a risk ladder that does not reflect accurately the typical magnitude of workplace risk. It is also possible to relate wages to a qualitative indicator of the perceived riskiness of the workplace environment, as researchers using the Quality of the Employment survey have done (Summers and Krueger, 1988; Biddle and Zarkin, 1988), but doing so does not produce an estimate of the VSL.

(4) $$p_i^* = p_i + e_i \quad ,$$

where *e* denotes the deviation between a worker's risk and the average fatality risk in his occupation or industry group. If e_i is uncorrelated with the true p_i and the other variables in the right-hand side of equation (3), equation (4) describes a classical errors-in-variables situation, which can be shown to result in a downward-biased $\hat{\beta}_2$, and hence in an understated VSL. Formally, the probability limit of the OLS coefficient on p_i^* is $\beta_2 \dfrac{Var(p)}{Var(p)+Var(e)}$, which is clearly less than β_2 (Greene, 2003, page 83-86).

To address this problem, Black and Kniesner (2003) suggest looking for two alternative measures of the same underlying risk. Let the first available measure be p_i^*, and the second be r_i^*, and assume that both are affected by classical measurement error. In other words, p_i^* is described by equation (4) above, and

(5) $$r_i^* = p_i + u_i \quad ,$$

where *u* is also a zero-mean measurement error. Further assume for the sake of simplicity that *e* and *u* are uncorrelated with one another, and that each of them is uncorrelated with the error term in equation (3). It can be easily shown that $\mathrm{Cov}(p^*, r^*) = \mathrm{Var}(p)$, i.e., the covariance between the two alternative measures of risks—both of which include measurement errors—is equal to the variance of true risk.

The quantity $\dfrac{Var(p)}{Var(p)+Var(e)}$ can, therefore, be estimated as the empirical covariance between p_i^* and r_i^*, the two observed measures of risk, divided by the empirical variance of p_i^*. The asymptotic bias of $\hat{\beta}_2$ is $\beta_2\left[\frac{Var(p)}{Var(p)+Var(e)}-1\right]$, so using data on p_i^* and r_i^* it is possible to calculate the term in brackets and get a sense of the extent to which the OLS estimates are underestimating the true β_2.[6]

Black and Kniesner (2003) illustrate this approach for the US, using wage and worker information from the 1995 Outgoing Rotation Groups of the Current Population Survey, and adopting mortality-specific industry rates as their p_i^* and occupation-specific mortality rates as their r_i^*. For all possible combinations of three-digit BLS and one-digit-by-state NIOSH fatality risk,[7] they find that $\text{Cov}(p^*, r^*)/\text{Var}(p^*)$ and $\text{Cov}(p^*, r^*)/\text{Var}(r^*)$ are very close to zero, which implies that the measurement error drowns out the "signal"—true risk. If the measurement error were truly classical, this would imply grossly understated estimates of the VSL. But because—depending on the specification of the regression equation—either the VSL estimates are very large or the OLS coefficient on the risk variables is negative, Black and Kniesner conclude that the measurement error is likely to be non-classical (i.e., correlated with some of the RHS variables in equation (3)).

To see why this might be the case, in some occupations there may be a great deal of heterogeneity in the actual job risk, and job assignment may be non-random in this respect. Consider, for example, the possibility that employers assign male and older convenience clerks to evening and late night shifts, when the risk of robberies is highest, reserving female and younger clerks to daytime hours. If we assign to each convenience

[6] If θ is an unknown parameter, the bias of estimator $\hat{\theta}$ is defined as $E(\hat{\theta})-\theta$.
[7] See section IV.A for a description of the BLS and NIOSH data.

store clerk the average risk in this occupation, we overestimate the risk for female and younger workers, and underestimate the risk for male and older workers.[8] In this case, the measurement error is correlated with other variables that appear in the right-hand side of the hedonic regression, resulting in a non-classical measurement error. Similar considerations apply for industry-based risk. Black and Kniesner conclude that even instrumental variable techniques are unlikely to remove the bias, and that the direction and magnitude of the bias of the OLS estimates is unknown.

B. Instability of Risk Coefficients

Even if one ignores the potential for measurement error, evidence from several studies points to the fact that the estimates of the VSL are not robust to even minor changes in the specification of equation (3), i.e., in the choice of the right-hand side variables. For example, in much empirical work industry dummies are entered in the right-hand side of (3) to capture inter-industry wage differentials, which have been widely documented to exist since Summers and Krueger (1988). Leigh (1995) finds that when both risk and industry dummies are included in the regression, the coefficient on risk is no longer significant. He interprets this to imply that inter-industry wage differentials, not compensating wages for risks of dying on the job, explain the positive correlations usually found between wages and mortality risks. He argues that workplace risks tend to be highly correlated with, and end up capturing, unpleasant aspects of jobs in certain industries. Another interpretation is, of course, that even broad industry

[8] In the presence of heterogeneity in the risk faced by workers, and if more highly skilled workers are capable of selecting into less risky jobs, Shogren and Stamland (2002) show theoretically that the VSL will be overestimated.

dummies are strongly correlated with workplace mortality risks, and that such collinearity makes it impossible to disentangle the effects of risks from those of one's industry.

In a recent study commissioned by the US Environmental Protection Agency, Black et al. (2003) explore the issue of stability of the estimates of the price of workplace risk, using three sources of data about individual workers and two sources of risk data. They show that (i) the estimates of the coefficient on risk vary dramatically with small changes in the inclusion of covariates in the right-hand side of the regression model; (ii) many of these coefficient estimates are negative, instead of positive; and (iii) using flexible functional forms confirms these OLS results.

C. Endogeneity or Unobserved Heterogeneity

There is reason to believe that job risk is endogenous with the wage rate, the dependent variable in the regression equation. This is because wage rates are likely to be affected by individual characteristics (e.g., skills) that are usually not well captured using the variables available in most datasets. These individual characteristics are likely to influence one's selection of job and job risks, resulting in correlation between risk and the error term in equation (3): $E(p_i\varepsilon_i) \neq 0$ and $E(q_i\varepsilon_i) \neq 0$.[9]

Yet, few empirical studies attempt to correct for endogeneity of wage and risk. Arabsheibani and Marin (2000) begin their study with treating risk as endogenous, which means that they must instrument for it, and at the same time allow for heterogeneous

[9] Using a very rich dataset (the national Longitudinal Study, which documents the Armed Forces Qualification Test scores), Black et al. find evidence that job risk is indeed correlated with individual characteristics and behaviors not usually available in most commonly used datasets. This means that in most compensating wage studies the error term in the regression model is correlated with the risk variable, making OLS estimates of the price of risk biased. Measurement error and self-selection are likely reasons why an effort to estimate a compensating wage model for Italian workers by Barone and Nese (2002) failed to detect a significant relationship between wage rates and objectively measured job risk.

preferences for risk and income by treating the coefficient on risk as a random variable. However, because of severe collinearity problems, they are not able to compute reliable VSL estimates when correcting for risk endogeneity and end up reporting only results based on treating risk as exogenous.[10] In a later study, Arabsheibani and Marin (2001) suggest that poor instruments are the main cause of these collinearity problems.

This implies that it should be possible to circumvent this problem by finding good instruments and/or by applying techniques based on panel data that construct instruments through transformations of whatever exogenous variables are available (such as the Hausman and Taylor approach, 1981). In the remainder of this paper, we apply the latter estimation technique and compare it with OLS and other instrumental-variable estimations.

IV. Application

We illustrate the abovementioned problems using data on worker characteristics from the British Household Panel Survey (BPHS). These data are collected through annual surveys, called waves, from 1991 through 2003. Each wave roughly corresponds to a different calendar year. Among other things, the dataset contains information about the workers' occupation, hours worked, earnings, experience, tenure with the present employer, education, and family status.

For waves 7, 11, 12, and 13 of the BHPS, which correspond to 1997, 2001, 2002, and 2003, we were able to match the BHPS data with workplace fatality rates for the worker's industry at the four digit-SIC code from the UK's Health and Safety Executive

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[10] When treating workplace mortality as exogenous, they find evidence of the existence of compensating wage differentials consistent with an average VSL to be around £9.7 million, with lower values for manual workers and larger values for managerial/professional workers.

(HSE). We express fatality rates as X per 100,000 a year. For waves 7 and 13, we also have mortality risk at the three-digit occupation level.

As in earlier studies, attention is restricted to full-time male workers aged 20-65 living in England and Wales. Because of the very distinct risk-related circumstances of their work environment, we excluded farmers, agricultural workers, firefighters, police officers, persons in the Armed Forces and security personnel from our sample.[11] We also created a subset of workers in blue-collar professions.[12] Descriptive statistics for the sample are reported in the Appendix.

A. Measurement Error in Our Data

Until recently, workplace fatality rates were computed at a relatively low level of resolution. For example, the majority of the US studies used two major sources of workplace fatality risk: the Bureau of Labor Statistics (BLS), which supplies total national counts of deaths by industry (up to the three-digit SIC code) or by occupation (at the three digit SOC level) for each year, and the National Institute of Occupational Safety and Health (NIOSH), which supplies rates at the one-digit industry or one-digit occupation for the different states.

Clearly, the latter has a low level of resolution, but the BLS data have many missing values, suppress death counts for cells with fewer than 5 deaths, do not account for geographic or seasonal variation in risk. Consequently converting the BLS counts

[11] SIC 92 categories A, B, P and Q and SOC 90 codes 600-619, 900-903.
[12] We define as blue-collar jobs SOC 80 codes 500-599 (craft and related occupations), 620-699 (personal occupations), 719-722 (sales assistants & check-out operators plus other sales representatives, 733 (scrap dealers and scrap metal merchants), 800-899 (plant & machine operatives, and 910-990 (other occupations except for agriculture, forestry & fishing).

into rates is fraught with difficulties.[13] Further, one might question the appropriateness of using industry-level risk rates, since this approach essentially treats a miner as facing the same risk as clerical personnel in the mining industry. A secretary, for example, would seem likely to face the same level of workplace risk whether he or she works for the primary metal industry sector or in a food processing establishment, suggesting that risk rates within occupation categories might be better measures of true risk.

Unfortunately, the potential for misclassifying one's occupation is high: Viscusi and Aldy (2003) report that earlier studies based on the 1977 CPS found that while 84% of workers surveyed and their employers agreed on industry affiliation, only 58% agreed on the worker's three-digit occupational status.[14] As a partial solution to this problem, Viscusi (2003) proposes using mortality rates at the two-digit industry and one-digit occupation code based on the recently released Census of Fatal Occupation Injuries. It is interesting that when he compares results with those based on a coarser aggregation of risk (two-digit industry), the VSL is actually higher in the latter case. Measurement error is presumably more substantial when industry-only risk figures are used, so if the measurement error is classical one would expect to the VSL to be lower, not higher, in the regression with by-industry-only data. By contrast, when by-occupation-only risk measures were used, the coefficient on risk was insignificant. When attention was

[13] One must divide death counts by the number of workers in the appropriate cells to form rates. Estimates of the number of persons employed in each industry or occupation are notoriously difficult to obtain. For example, Viscusi (2003) uses the Current Population Survey (CPS) to produce estimates of the number of workers in each industry-occupation cell, but this approach is criticized by Taylor (2004), who points out that the CPS is not representative of occupations or firms, and uses instead the Occupation-Employment Matrix (OEM). In both cases, an additional source of measurement error is the sampling variation associated with very small cell sizes.

[14] Also see Bound et al (2001).

restricted to blue-collar workers, a substantially higher VSL resulted, which is against expectations.

Does using risk data with finer resolution help reduce measurement error problems? To answer this question, we turn to our data from the BHPS and the HSE, and compute the covariance between mortality rates at the four-digit industry level (our p_i^*) and mortality rates at the three-digit occupation level (our r_i^*) as suggested by Black and Kniesner (2003) (see section III.A) for 1997 and 2003, the only two years for which both sets of risk data are available. Our calculations are reported in table 1.

Table 1. Variances of mortality risk figures and true mortality rates

	All workers		**Blue collar workers**	
Year	**1997**	**2003**	**1997**	**2003**
Nuber of obs.	1895	3289	1146	1299
Mean fatsic	1.73949	1.069444	2.244604	1.680139
Mean fatsoc	1.818226	1.224681	2.783882	2.139557
Var(fatsic)	17.21954	6.185134	17.25135	8.299913
Var(fatsoc)	19.73716	2.921962	29.38772	3.797896
Covar(fatsic, fatsoc)	0.3749	0.339	0.4544	0.3136
[A] Cov(fatsic, fatsoc)/Var(fatsoc)	0.021771778	0.054808837	0.026339968	0.037783529
[B] Cov(fatsic, fatsoc)/Var(fatsoc)	0.0189946	0.1160179	0.0154622	0.082572

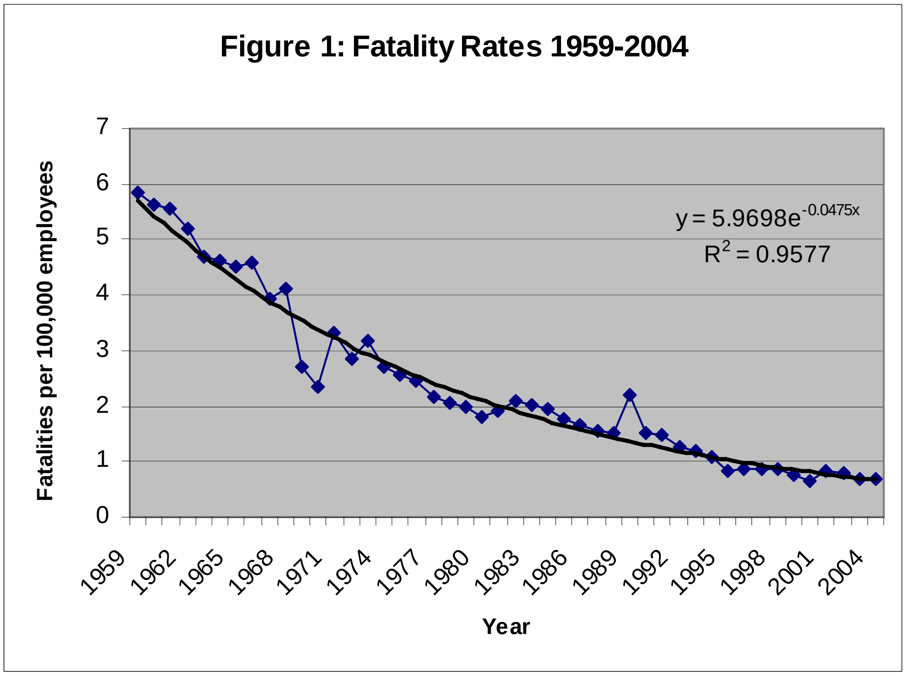
Legend: fatsoc=fatality rate by occupation; fatsic=fatality rate by industry.

We remind the reader that when the fatality risk is affected by a classical measurement error, then the probability limit of the OLS estimates of the coefficient on risk is $\hat{\beta}_2[Var(p)/(Var(p)+Var(e))]$, where p is true risk and e is the measurement error. The figures in the last two rows of table 1 are precisely the term in brackets in this expression when (i) the regressor in the wage equation p_i^* is risk measured at the industry level (row [A]) and (ii) the regressor is risk r_i^* measured at the occupation level, which means that r replaces p in the denominator of the expression for the probability limit of $\hat{\beta}_2$ above (row [B]). Clearly, both sets of figures are very close to zero, implying that if the measurement error were truly classical, one would expect the estimates of the VSL based on OLS regressions to be grossly downward biased. Both sets of ratios are similar to—and even lower than—those computed by Black and Kniesner for the US, despite the greater detail of our fatality rate data.

We conclude that even with our 4-digit-level risk data the measurement error seems to be drowning out the "signal," i.e., the true mortality risk. It would be interesting to find out whether the pervasiveness of the measurement error has increased over the years as mortality rates have declined, which would make it increasingly difficult to disentangle compensating wage differentials from other factors as one uses more recent data. As shown in Figure 2, which refers to total workplace mortality rates, in the UK fatal workplace accident rates per 100,000 employees fell from about 5.8 in 1959 to 0.7 in 2004. The decline has been quite steady, approximately following an exponential decrease rate of 4.75% per annum. The shape of the curve suggests that it is increasingly difficult to decrease workplace accident risk even further.[15] Unfortunately, the period over which we were able to find mortality rate data at a fine level of resolution is too short and too recent to confirm or refute our conjecture about the error term becoming an increasingly important component of workplace fatality attributed to workers in various industries and occupations.

[15] Fatality rates declined by 3.1 per 100,000 in the ten years between 1959 and 1969, but the corresponding decrease between 1994 and 2004 was only 0.36. If this process continues at the same rate, only 0.28 lives per 100,000 employees will be saved by safety measures imposed between 2004 and 2014. Further, workplace accidents account for a smaller and smaller share of all-cause mortality risks. For males aged 25-35, for example, the ratio of job mortality risk to the risk of dying for all causes fell from 2.7 % in 1976 to 0.7 % in 2003, the year of wave 13 of the BHPS. Similar trends are seen for other age groups as well.

Figure 2: UK overall fatality rates 1959-2004. All industries and occupations



B. Robustness of VSL Estimates to the Inclusion/Exclusion of Variables

As discussed above, earlier literature has suggested that the compensating wage differentials, and hence the estimates of the VSL, are not robust to adding explanatory variables—such as industry dummies or non-fatal injury risk—to the right-hand side of the wage equation. Are these results common when fitting hedonic wage equations, or are they specific to the US data for which they were originally claimed?

To answer this question, we specify a linear model similar to equation (3), where the dependent variable is the logarithm of annual wages and the independent variables include education, experience, experience squared, overtime worked, tenure with the present employer, a good health dummy, a race dummy, union status, regional dummies, plus the risk variable(s) and, depending on the specification, industry and occupation dummies. The model is estimated using OLS.

The results of this exercise are shown in table 2. They suggest that for the full sample the VSL remains relatively stable when industry and occupation dummies are added to the regression equations (columns A and B, respectively), but that further controlling for non-fatal risk almost triples the VSL (column C), which then falls by almost one-half when industry and occupation dummies are further added.[16]

The second panel shows that the VSL is generally much higher among blue collar workers. Caution should be used when interpreting these results. Viscusi (2003) notes that if most accidents happen to blue-collar workers, but fatality rates are computed using all workers in a particular industry as the denominator, blue-collar risks will be understated and their VSL overstated. Since we are unable to calculate a risk rate specific for blue collar workers only, our estimates of the VSL for blue collar workers are potentially affected by this problem. This problem should disappear when one uses risk by occupation, but this comes at the expense of misclassification and other problems.

The VSL for blue-collar workers changes dramatically when occupation and industry dummies are added in the basic specification, which uses fatal risk only (columns A and B). However, when one controls for fatal and non-fatal risk from the start, the VSL in the no-dummy specification (column C) is similar to that of column A, and adding further industry and occupation controls has a relatively modest effect (a 15% decline) on the VSL.[17]

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[16] The VSL is calculated as the product of β_2 by average wage, times 100,000 (because risk is expressed as XE-05).

[17] As Leigh (1995) has shown, apparent wage/risk differentials can be caused by inter-industry wage differentials. In our data, coefficients remained significant even after including industry dummies, which is

Table 2. OLS coefficient on risk variables for different model specifications

Risk based on SIC, OLS	**Fatal risk only**		**Fatal and non-fatal risk**	
	A	**B**	**C**	**D**
	No industry or occupation dummies	**With industry & occupation dummies**	**No industry or occupation dummies**	**With industry & occupation dummies**
All workers (n=5638)				
Coeff.	0.0022504	0.0023359	0.0065263	0.0038613
t-stat	1.21	1.23	2.71	1.67
Coeff. on nonfatal			-0.0001140	-0.0000448
t-stat			-2.80	-1.15
VSL (million 1996 £)	4.63	4.80	13.42	7.94
Blue collar workers (n=2559)				
Coeff.	0.0116428	0.0071250	0.011388	0.0096666
t-stat	5.64	3.12	4.34	3.50
Coeff. on nonfatal			0.0000083	-0.0000881
t-stat			0.16	-1.63
VSL (million 1996 £)	18.91	11.57	18.50	15.70

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

probably due to the fact that risk is measured on the 4-digit SIC level, whereas the industry dummies were based on major categories.

C. Exploiting the Panel Structure of the Data

To our knowledge, this is the first time a panel dataset is used for the purpose of estimating the price of risk implicit in labor markets. We therefore have the opportunity to control explicitly for all omitted determinants of wage, even when they are correlated with risk. This is usually done by specifying fixed effects (FE) model, i.e., models with individual-specific intercepts, and fitting the so-called "within" estimator, or, alternatively, by adopting random effects (RE) models. The latter use a common intercept for all cross-sectional units, but the error term is broken down into two components, one of which remains fixed over time within one individual (but varies across individuals).

Random effects models are estimated using GLS, which yield consistent and efficient estimates as long as both components of the error term are uncorrelated with the regressors. By contrast, the "within" estimator is unbiased, even if the unobserved individual-specific effects are correlated with included regressors. Large differences between the within and GLS estimates of the coefficient on risk would be taken as evidence that the worker-specific effects *are* correlated with the included regressors.[18]

Results for the FE and the RE models are reported in the top panel of table 3 for all workers and of table 4 for blue-collar workers. For the full sample, the "within" and GLS coefficients on fatal risk are remarkably similar to one another, and they are *negative* and significant. They imply that people are, if anything, paid *less* when they take higher-risk jobs, refuting the very notion of compensating wage differentials. The coefficient on nonfatal risk changes more dramatically across "within" and GLS

[18] This is done formally using a Hausman test (see Hsiao, 2003).

estimation procedure, but in both cases it is not statistically significant at conventional levels.

Blue collar workers are a completely different story: The risk coefficients vary widely across the model specification and estimation techniques, but are always statistically insignificant. Ignoring for the moment the issue of coefficient significance, both the "within" estimator and GLS (for fixed and random effects models, respectively) produce reasonable VSL figures for blue-collar workers (2-6 million 1996 £ or 3.3-10 million 2000 $). Although both are within the normally accepted range, there is a three-fold difference between them.

D. Endogeneity of Risk

As noted earlier, even if one ignores the potential for measurement error, it is reasonable to suspect that workplace risk and wages are endogenous. Unobserved individual characteristics that influence wages but get soaked up in the error term of (3) may well be correlated with workplace risk, making the OLS estimates biased and inconsistent.

Various approaches are possible to address endogeneity. One of them is to specify fixed effects model and estimate them using the "within" estimator, as in the previous section. Alternatively, one can instrument for workplace risk and estimate equation (3) using 2SLS. We do indeed apply 2SLS. Our instruments are all of the right-hand side variables of equation (3) plus—to ensure identification—the worker's marital status, his nonlabor income, wife's income, number of dependent children, home ownership status, and dummies for his father's social class (for summary statistics see Table A1 in the Appendix).

A further variant of 2SLS is deployed by Arabsheibani and Marin (2000, 2001), who treat risk as endogenous and preferences for risk and income as heterogeneous by allowing the coefficient on risk to be random. Formally, fatal and non-fatal risks are expressed as

(6) $p_i = \gamma_0 + \mathbf{z}_i\gamma_1 + \varepsilon_{2i}$

(7) $q_i = \delta_0 + \mathbf{z}_i\delta_1 + \varepsilon_{3i}$.

In equations (6) and (7), the vector $\mathbf{z}_i$ includes all covariates contained in $\mathbf{x}_i$ plus, as per the definition of instrument, additional variables that determine risk choice but are not correlated with the error term in the wage equation. We chose the same instruments as for our basic 2SLS procedure.

Coefficients β_2 and β_3 in equation (3) are replaced with $\beta_{2i} = \overline{\beta}_2 + u_{1i}$ and $\beta_{3i} = \overline{\beta}_3 + u_{2i}$, respectively. After substituting these into the wage equation, we obtain:

(8) $$\ln w_i = \beta_0 + \mathbf{x}_i\beta_1 + p_i\overline{\beta}_2 + p_iu_{1i} + q_i\overline{\beta}_3 + q_iu_{2i} + \varepsilon_{1i}$$

Garen (1984, 1988) proposed a consistent two-step estimation procedure for this model. In the first step, one runs OLS on equations (6) and (7) and forms the residuals $\hat{\varepsilon}_{2i}$ and $\hat{\varepsilon}_{3i}$. In the second step, one runs OLS on the equation:

(9) $$\ln w_i = \beta_0 + \mathbf{x}_i\beta_1 + p_i\overline{\beta}_2 + c_1\hat{\varepsilon}_{2i} + c_2\hat{\varepsilon}_{3i} + c_3\hat{\varepsilon}_{2i}p_i + c_4\hat{\varepsilon}_{2i}q_i + c_5\hat{\varepsilon}_{3i}p_i + c_6\hat{\varepsilon}_{3i}q_i + e_i$$

The results from 2SLS and the Garen procedure are displayed in the second panel of table 3 (for the full sample) and table 4 (for blue-collar workers). For simplicity, we restrict attention to the models that include only fatal risk. Again, for the full sample of workers the risk coefficient is always negative, refuting the notion of compensating wage differentials for risk. By contrast, with the blue-collar subsample, these two instrumental-variable procedures results in large coefficients on risk and correspondingly large VSLs ranging from 70 to 102 million pounds sterling (116 to 170 million 2000 dollars).

Table 3: Compensating wage regression results for the full sample

Risk based on SIC, all workers	**Fatal risk only**		**Fatal and nonfatal risk**	
	No industry or occupation dummies	**With industry & occupation dummies**	**No industry or occupation dummies**	**With industry & occupation dummies**
FE (n=4796)				
Coeff. on mortality rates	-0.0030549	-0.0029672	-0.0026492	-0.0025267
t-stat	-2.00	-1.84	-1.38	-1.29
Coeff. on nonfatal injury rates			-0.0000115	-0.0000132
t-stat			-0.35	-0.39
VSL	n/a	n/a	n/a	n/a
RE (n=4796)				
Coeff. on mortality rates	-0.0033346	-0.0036314	-0.0018903	-0.0029294
z-stat	-2.23	-2.24	-1.00	-1.48
Coeff. on nonfatal injury rates			-0.0000403	-0.0000208
t-stat			-1.26	-0.62
VSL	n/a	n/a	n/a	n/a
2SLS (n=5638)				
Coeff. on mortality rates	-0.0357515	-0.0221513		
t-stat	-2.18	-1.08		
VSL	n/a	n/a		
Garen proc. (n=5638)				
Coeff. on mortality rates	-0.0347135	0.0049412		
t-stat	-2.19	0.34		
VSL	n/a	10.16		
Hausman-Taylor (n=4796)				
Coeff. on mortality rates	-0.0030382	-0.0031122	-0.0027392	-0.0028191
z-stat	-1.89	-1.97	-1.36	-1.46
Coeff. on nonfatal injury rates			-0.00000851	-0.0000088
t-stat			-0.25	-0.27
VSL	n/a	n/a	n/a	n/a

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

Table 4: Compensating wage regression results for blue-collar workers

Risk based on SIC, blue collar workers only	Fatal risk only		Fatal and nonfatal risk	
	No industry or occupation dummies	**With industry & occupation dummies**	**No industry or occupation dummies**	**With industry & occupation dummies**
FE (n=1974)				
Coeff. on mortality rates	0.00000139	0.0009226	0.0011376	0.0016023
t-stat	0.00	0.40	0.44	0.60
Coeff. on nonfatal injury rates			-0.0000411	-0.0000271
t-stat			-0.79	-0.50
VSL	0.002	0.001	1.85	2.60
RE (n=1974)				
Coeff. on mortality rates	0.0028377	0.0028991	0.0037008	0.0038576
z-stat	1.36	1.29	1.46	1.46
Coeff. on nonfatal injury rates			-0.0000301	-0.000036
t-stat			-0.60	-0.69
VSL	4.61	4.71	6.01	6.27
2SLS (n=2559)				
Coeff. on mortality rates	0.0596566	0.0432909		
t-stat	4.22	2.47		
VSL	96.89	70.31		
Garen proc. (n=2559)				
Coeff. on mortality rates	0.0625513	0.0537317		
t-stat	4.89	4.21		
VSL	101.60	87.27		
Hausman-Taylor (n=1974)				
Coeff. on mortality rates	0.0000812	0.0008241	0.0007656	0.001222
z-stat	0.04	0.35	0.31	0.44
Coeff. on nonfatal injury rates			-0.0000247	-0.0000158
t-stat			-0.51	-0.28
VSL	0.13	1.34	1.24	1.98

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

These results are not credible, and we believe that the explanation is that the instruments we have used do not explain much of the variation in p_i, something that the Garen procedure seems vulnerable to. Only roughly 3-5% of the variation in job risk for all workers can be predicted using the instruments, with the remainder being absorbed into the residuals $\hat{\varepsilon}_{2i}$. As a consequence, the residuals are highly correlated with job risk and with $\hat{\varepsilon}_{2i} p_i$ in equation (9). The correlation coefficient between $\hat{\varepsilon}_{2i}$ and p_i is 0.98 for all workers and also for blue-collar workers, which explains why the regression coefficient β_2 is so sensitive to the inclusion of the residuals.[19] Arabsheibani and Marin (2001) report that they encountered the very same problem, despite using a broader set of instruments and obtaining much better first-stage R squares, and conclude that compensating risk differentials extracted in this way should be viewed with great caution due to multicollinearity.

Given how difficult it is to find suitable instruments for risk, is it possible to exploit the panel structure of the data and an estimation technique developed by Hausman and Taylor (1981) to remedy the problem of correlation between unobserved characteristics of the worker and risk? Consider the regression equation:

$$(10) \qquad \mathbf{y}_i = \mathbf{X}_i\beta + \mathbf{Z}_i\alpha + \left[v_i\iota + \eta_i\right] \quad ,$$

where $\mathbf{y}_i$ is a T×1 vector of observations on the dependent variable for individual *i*, $\mathbf{X}_i$ is a T×k matrix of time-varying regressors and $\mathbf{Z}_i$ is a T×G matrix of time-invariant regressors. In the term in brackets, v_i is the individual-specific component of the error

[19] It should also be noted that the coefficient on risk variable is of the opposite sign of that on the residuals.

term, ι is a T×1 vector of ones, and η_i is a T×1 vector of independent error terms. After stacking individual vectors and matrices of variables, equation (10) becomes

$$(11) \qquad \mathbf{y} = \mathbf{X}\beta + \mathbf{Z}\alpha + [\mathbf{VW} + \varsigma] \ ,$$

where **y**, **X**, **Z** and ς have nT rows, **V** is an nT×1 matrix of individual dummies, and **W** is an n×1 vector of idiosyncratic error terms. Further let $\boldsymbol{\Omega}$ denote the variance covariance matrix of the error terms in brackets.

The Hausman-Taylor approach first distinguishes between exogenous and endogenous (i.e., correlated with **W**) time-varying and time-invariant regressors, namely $\mathbf{X} = [\mathbf{X}_1 \quad \mathbf{X}_2]$, and $\mathbf{Z} = [\mathbf{Z}_1 \quad \mathbf{Z}_2]$, where the subscripts 1 and 2 denote the exogenous and endogenous subsets, respectively. The Hausman-Taylor is an instrumental-variable approach corrected for the non-spherical nature of the variance-covariance matrix of the error terms in the brackets in equation (11):

$$(12) \qquad \begin{bmatrix} \hat{\beta} \\ \hat{\alpha} \end{bmatrix} = \left[\mathbf{W}'\Omega^{-1/2}\mathbf{P_A}\Omega^{-1/2}\mathbf{W}\right]^{-1}\mathbf{W}'\Omega^{-1/2}\mathbf{P_A}\Omega^{-1/2}\mathbf{y} \,,$$

where $\mathbf{P_A} = \mathbf{A(A'A)^{-1}A'}$, and **A** is a matrix of instruments. Hausman and Taylor propose using (i) the deviations from the individual's means of the $\mathbf{X}_1$, (ii) the deviations from the means of $\mathbf{Z}_1$, (iii) the $\mathbf{X}_1$ and the $\mathbf{Z}_1$ as instruments.

We report the results of the Hausman-Taylor procedure in the second panel of tables 3 and 4 for all workers and the blue-collar sample. We treat education as an endogenous time-invariant regressor, and experience, experience squared, tenure, union status, and occupation dummies as endogenous time-varying regressors. Race, good health, geographical dummies, job overtime, and industry dummies are regarded as exogenous variables.

The results show that whether or not one controls for non-fatal risk, the coefficients on fatal risk are *negative*, but negligible and statistically insignificant for the full sample. For blue-collar workers, the risk coefficients are positive, but still insignificant. They imply (ignoring for the moment the issue of significance) a VSL of £1.24 to 1.98 million—an order of magnitude from the values we obtained using 2SLS and the Garen procedure.[20]

V. Conclusions

Unlike previous studies, we use panel data documenting wages and other individual characteristics of UK workers to examine whether there is evidence of compensating wage differentials for workplace risk, and to recover estimates of the VSL implicit in workers' job choices. We have risk data at the four-digit industry level and three-digit occupation level.

If occupation- and industry-level risks are both assumed to be observing true risk with measurement error, and this measurement error is assumed to be classical (i.e., uncorrelated with individual characteristics in the right-hand side of the wage equation), the measurement error in our data drowns out the true risk "signal." This problem may have worsened over the years as observed workplace mortality risks have declined.

Under the assumption that the measurement error is classical, a substantial measurement error relative to the true signal would imply that the OLS coefficient on risk

[20] Results are qualitatively similar when we use risk data by occupation (see tables A.2-A.4 in the Appendix).

grossly underestimates the true association between risk and wages. But this expectation is at odds with the OLS estimates of the VSL we obtain for the full sample and for blue-collar workers: Both are large and for the most part sensitive to specification changes.

We also employ estimation procedures that correct for the endogeneity of risk and wages. When we do, the coefficient on risk is either negative (for the full sample) or, as in the case of blue-collar workers, disproportionately large, to the point that one would infer the VSL to be in the range of 70 to 102 million 1996 £ (116 to 170 million 2000 US $). These results raise doubts about the existence and/or credibility of compensating wage differentials. Possible explanations for these findings are that we do not have good instruments for workplace risk, and that our risk data most likely understate the risk for blue-collar workers (but are more reliable when used for all workers).

Our panel data offer us the unique opportunity to control for unobserved worker characteristics that influence both wages and workplace risk. Fixed effects models, which we estimate using the "within" estimator, allow the individual-specific effects to be correlated with included worker characteristics. In this case, we find no evidence of compensating wage differentials. We also use the Hausman-Taylor estimation approach in hopes of increasing efficiency while allowing for the endogeneity of risk and other covariates, and again find no significant association between wages and risk. Both the "within" estimator and the Hausman-Taylor approach find, if anything, a *negative*, usually insignificant, association between wages and workplace risk for the full sample—against the very notion of compensating wage differentials.

When attention is restricted to blue-collar workers, only the estimation approaches based on panel data produce estimates of the VSL (1.2-2.6 million 1996 £ or

2.0-4.3 million 2000 US $) that are within the range deemed appropriate for this construct (Viscusi and Aldy, 2003), but the coefficient on fatal risks is not statistically significant. Clearly, we need to interpret this finding with caution, since the risk measure is likely to overstate the true risk of blue-collar workers. Yet it would seem that when we use estimation techniques based on panel data the results are much more reasonable than we use estimation techniques that treat our data as independent cross sections. This raises concern about most of the existing estimates of the VSL, which are indeed based on cross-sectional samples.

We also note from the data that workplace fatality risk has been declining sharply over the period 1959-2004. This reflects changes may be legally and technologically driven, making the notion of a stable value for VSL that can be applied in a policy context very suspect. Such a reduction in the demand for risk on the part of employers legally mandated to reduce risk level should lower the equilibrium wage premium.

In practice, however, we have not been able to track the dynamics of such a shift, although this would make for interesting further work on the subject. We conclude that *if* compensating differentials for risk exist, measurement error and other econometric problems, as well as a changing workplace risk environment prevent us from observing them.

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Appendix.

Table A1: Descriptive statistics

Waves 7, 11-13	All workers (n=5638)		Blue collar (n=2559)	
Variable	**Mean**	**St. Dev.**	**Mean**	**St. Dev.**
Labor income and job risk variables				
incdef	20556	12288	16241	6920
avfatal[a]	1.276	3.047	1.866	3.338
avmajorinj[a]	161.5	181.8	204.9	166.8
Continuous variables in X				
educ	10.30	1.11	9.85	0.83
experience	24.53	10.72	24.61	11.14
expsq	716.93	565.47	729.92	591.12
jbot	5.63	7.12	5.93	7.40
tenure	5.38	6.60	6.24	7.21
Dummy variables in X				
healthy	0.923	0.266	0.915	0.279
white	0.962	0.192	0.971	0.169
union	0.318	0.466	0.371	0.483
inlondon	0.022	0.147	0.011	0.106
outlondon	0.053	0.225	0.030	0.172
southeast	0.186	0.389	0.143	0.351
swea	0.132	0.339	0.147	0.354
midlands	0.173	0.379	0.184	0.388
manmersey	0.049	0.217	0.042	0.200
wales	0.194	0.395	0.229	0.420
y1997	0.259	0.438	0.264	0.441
y2001	0.298	0.457	0.294	0.456
y2002	0.210	0.407	0.208	0.406
Occupation dummies				
managers	0.208	0.406	0.000	0.000
professionals	0.116	0.320	0.000	0.000
associates	0.101	0.301	0.000	0.000
clerical	0.093	0.291	0.000	0.000
craft	0.187	0.390	0.412	0.492
service	0.024	0.152	0.052	0.223
sales	0.041	0.199	0.036	0.187
operatives	0.172	0.377	0.379	0.485
Industry dummies				
mining	0.005	0.070	0.009	0.096
manufacturing	0.315	0.465	0.418	0.493
construction	0.079	0.270	0.122	0.328
wholesale	0.126	0.332	0.127	0.333
hotelrest	0.018	0.134	0.025	0.156
transport	0.103	0.305	0.143	0.350
realestate	0.115	0.320	0.045	0.207
public	0.077	0.266	0.024	0.154
health	0.035	0.183	0.022	0.146
social	0.038	0.192	0.041	0.197
finance	0.026	0.158	0.003	0.052
education	0.048	0.214	0.006	0.079
Additional instruments in Z				
nlincdef	794	2670	576	1698
nkids	0.752	1.011	0.792	1.056
workspouse	0.643	0.479	0.617	0.486
married	0.651	0.477	0.635	0.482
pasec 1 to 31	Series of dummies for father's social class			

a: Based on 4-digit SIC. Values based on 3-digit SOC are: All workers (n=2456): 1.351 (fatal), 121.5 (nonfatal); blue collar workers (n=1266): 2.266 (fatal), 197.9 (nonfatal).

Table A2. OLS coefficient on risk variables for different model specifications, by SOC

Risk based on SOC, OLS	**Fatal risk only**		**Fatal and non-fatal risk**	
	A	**B**	**C**	**D**
	No industry or occupation dummies	**With industry & occupation dummies**	**No industry or occupation dummies**	**With industry & occupation dummies**
All workers (n=2456)				
Coeff.	-0.0122041	-0.0010326	0.0017806	0.0008213
t-stat	-3.89	-0.33	0.45	0.22
Coeff. on nonfatal			-0.0003556	-0.0000546
t-stat			-5.76	-0.91
VSL (million 1996 £)	n/a	n/a	3.51	1.62
Blue collar workers (n=1266)				
Coeff.	0.0037013	-0.0005689	0.0032278	0.0002676
t-stat	1.24	-0.18	0.89	0.07
Coeff. on nonfatal			0.000014	-0.00000782
t-stat			0.23	-0.13
VSL (million 1996 £)	5.81	n/a	5.07	0.04

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

Table A3: Compensating wage regression results for the full sample, by SOC

Risk based on SOC, all workers	**Fatal risk only**		**Fatal and nonfatal risk**	
	No industry or occupation dummies	**With industry & occupation dummies**	**No industry or occupation dummies**	**With industry & occupation dummies**
FE (n=1158)				
Coeff. on mortality rates	-0.0022404	-0.00345	-0.0032273	-0.004424
t-stat	-0.54	-0.8	-0.67	-0.88
Coeff. on nonfatal injury rates			0.0000277	0.0000272
t-stat			0.41	0.39
VSL	n/a	n/a	n/a	n/a
RE (n=1158)				
Coeff. on mortality rates	-0.008789	-0.0062546	-0.0069949	-0.007733
z-stat	-2.36	-1.62	-1.58	-1.71
Coeff. on nonfatal injury rates			-0.0000505	0.0000428
t-stat			-0.78	0.65
VSL	n/a	n/a	n/a	n/a
2SLS (n=2505)				
Coeff. on mortality rates	-0.1079389	-0.0450135		
t-stat	-3.99	-1.69		
VSL	n/a	n/a		
Garen proc. (n=2505)				
Coeff. on mortality rates	-0.1132858	-0.0568773		
t-stat	-4.93	-2.74		
VSL	n/a	n/a		
Hausman-Taylor (n=1158)				
Coeff. on mortality rates	-0.0021184	-0.00348	-0.003156	-0.0043135
z-stat	-0.61	-0.91	-0.78	-0.98
Coeff. on nonfatal injury rates			0.0000292	0.00002
t-stat			0.51	0.36
VSL	n/a	n/a	n/a	n/a

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

Table A4: Compensating wage regression results for blue-collar workers, by SOC

Risk based on SOC, blue collar workers only	Fatal risk only		Fatal and nonfatal risk	
	No industry or occupation dummies	With industry & occupation dummies	No industry or occupation dummies	With industry & occupation dummies
FE (n=1974)				
Coeff. on mortality rates	-0.0035649	-0.0036039	-0.0042803	-0.0039749
t-stat	-0.8	-0.76	-0.86	-0.73
Coeff. on nonfatal injury rates			0.0000219	0.00000876
t-stat			0.32	0.12
VSL	n/a	n/a	n/a	n/a
RE (n=1974)				
Coeff. on mortality rates	-0.004751	-0.0076434	-0.0048982	-0.00813
z-stat	-1.21	-1.82	-1.09	-1.69
Coeff. on nonfatal injury rates			0.0000134	0.0000163
t-stat			0.2	0.24
VSL	n/a	n/a	n/a	n/a
2SLS (n=2559)				
Coeff. on mortality rates	0.0108529	-0.0097568		
t-stat	0.65	-0.54		
VSL	17.04	n/a		
Garen proc. (n=2559)				
Coeff. on mortality rates	0.0096179	0.0081068		
t-stat	0.56	0.49		
VSL	15.1	12.73		
Hausman-Taylor (n=1974)				
Coeff. on mortality rates	-0.0034855	-0.0036715	-0.004205	-0.0038965
z-stat	-0.80	-0.97	-0.89	-0.89
Coeff. on nonfatal injury rates			0.0000213	0.000006
t-stat			0.33	0.11
VSL	n/a	n/a	n/a	n/a

All regressions control for experience, experience squared, overtime hours worked, tenure, and dummies for health status, race and location. The VSL is expressed in million British pounds in 1996 prices; to convert to 2000 dollars, multiply by 1.66.

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