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Valuing risk reductions: Incorporating risk heterogeneity into a revealed preference framework

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ABSTRACT

Our study expands the hedonic wage framework to take advantage of the inherent differences in workplace deaths, both in type and probability of occurrence, and examines revealed preferences over these heterogeneous risks. We use data on all fatal workplace deaths in the US from 1992 to 1997 and develop risk rates that are differentiated by how the fatal injury occurred. Within sample tests of the equality of compensating wage differentials for heterogeneous risks indicate that we can reject aggregation of homicide risks with other sources of workplace fatalities. However, our results are not without qualification and highlight important nuances of the labor market as related to estimating compensating wage differentials for risks that have generally been ignored in the previous literature.

"All deaths are bad. But some deaths seem worse than others." Cass R. Sunstein (p. 259, 1997)

1. Introduction

Benefit/cost evaluations of programs affecting mortality among the public requires estimating a monetary benefit from reducing the probability of death—the so-called "value of a statistical life (VSL)." In an attempt to develop such estimates, a large empirical literature has developed analyzing the tradeoffs individuals make when accepting more or less risk in their lives (for recent reviews, see McConnell, 2006; Kochi et al., 2006; Viscusi and Aldy, 2003; Mrozek and Taylor, 2002). Yet, as US Senate Bill S.3564 indicates, the use of VSL estimates in policy analysis continues to generate considerable controversy (United States S. 3494, 2008). This bill is a reaction to recent publicity

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regarding the use of VSL estimates in environmental policy analyses (Borenstein, 2008; Fahrenthold, 2008), and if enacted would limit the benefits analyses of the US Environmental Protection Agency to the highest value used in any previous analysis and would only allow upward adjustments to account for income growth and inflation. The bill also rejects attempts to differentiate the VSL by the characteristics of the population at risk, and importantly, rejects the implied benefits transfer in using VSL estimates based on labor market studies to value environmental risks. Specifically, the bill states that "(4)(A) there is a great difference between a voluntarily accepted risk and an involuntarily imposed risk; and (B) that difference renders the use of a value of statistical life based on measures of voluntarily accepted risks questionable as applied to involuntarily imposed risks;" (United States S. 3494, 2008). This is an explicit recognition that if the character of a risk affects the valuation of reducing the risk, policy-relevant VSL estimates should be individuated by the type of threat targeted in the policy.

There has long been the recognition that individual perception and assessment of risks, and attitudes towards risk, vary by the type of threat (see Slovic, 2000; Lichtenstein and Slovic, 2006 for recent summaries of this literature).[1] Perceived characteristics of a threat such as the amount of dread it evokes, the level of controllability one has in the face of it, and how observable or well-known it is to those affected and to science, are among factors shown to bias an individual's risk perceptions (Slovic et al., 1980). This has motivated studies to examine how characteristics of a threat affect the value placed on risk reduction. Early evidence in these regards was presented by Jones-Lee et al. (1985) who conducted a national survey eliciting willingness to pay (WTP) for risk reductions over three types of threats: cancer, heart disease and traffic accidents. They find significant differences in WTP among the threats and report "that people do make significant distinctions between different ways of dying" (p. 68). Savage (1993) and Hammitt and Liu (2004) also find evidence that hazards which are characterized by dread and/or viewed as uncontrollable result in higher WTP estimates to reduce the risk.[2] These results are consistent with the literature on risk assessment which indicates that threats are considered more severe which are perceived as highly dreaded and uncontrollable (Slovic et al., 1980, 1981; Sunstein, 1997; Slovic and Peters, 2006).

While the existing literature indicates that characteristics of the event influence risk valuation, the empirical evidence is limited to surveys. We have no revealed preference evidence regarding heterogeneous risk valuation. Past studies estimating compensating wage equations use data on the count of total fatal occupational injuries, usually in conjunction with data on total non-fatal occupational injuries, and treat all workplace deaths as equivalent so the risk of death is undifferentiated by the cause of death. This assumes people view the outcome, death, as the only factor in the wage negotiation and their utility is not affected by the event leading to the fatality. Yet workplace deaths are not homogeneous events. While traditional workplace accidents such as being struck by machinery, falls, and motor-vehicle related events still comprise the majority of workplace accidents that lead to fatalities, since the 1970s the reported incidence of homicide in the American workplace has grown substantially and accounted for approximately 20% of workplace deaths by the mid-1990s. Slovic et al. (1980) specifically identify crime-related events as rating highly dreaded and uncontrollable, while incidents related to motor-vehicle and power-tool accidents are rated the opposite. As such, we would expect that compensating wage differentials could be significantly different for facing homicide risks as compared to other, more traditional, workplace hazards.

Our study expands the hedonic wage framework to take advantage of the inherent differences in workplace deaths, both in type and probability of occurrence, and examines revealed preferences over these heterogeneous risks. We use data on all fatal workplace deaths in the US from 1992 to 1997 and develop risk rates that are differentiated by how the fatal injury occurred. We focus on three primary sources of workplace fatalities – transportation-related fatalities, other traditional hazards (such as contact with heavy equipment or falls), and homicide – and conduct within sample tests of the equality of compensating wage differentials for these heterogeneous risks. Results indicate that we

[1] We use the terms perception, assessment and attitude interchangeably here; our objective is to distinguish the probability, or risk, of an event, with people's judgement or concern about the event leading to or causing the risk.

[2] See also Breyer and Felder (2005), Chilton et al. (2006) and Treich (2008) for recent examples of surveys that consider how characteristics of risk, including the baseline risk, affect risk valuation.

can reject aggregation of homicide with other sources of workplace fatalities, a finding with implications beyond cost/benefit analysis of labor market safety policies. If we can reject risk aggregation within our relatively narrow set of events, then transferring VSL estimates from labor market studies to policy contexts involving very different events is certainly questionable. The US EPA applied a mean VSL estimate of $7.5 million (in 2008 dollars) to assess regulatory program benefits due to the CleanAir Interstate Rule (US EPA, 2005). This VSL estimate is based primarily on labor market studies of instantaneous workplace deaths. The extent to which a VSL estimate based on instantaneous workplace deaths can be used to evaluate policy benefits from reducing the incidence of cancer, respiratory illness-induced mortality, or other dissimilar events associated with air pollution remains to be seen.

Our empirical approach and findings are not without qualification however, and highlight important conceptual and empirical issues associated with the use of hedonic wage models for identifying willingness to pay for risk reductions. Conceptually, we employ an expected utility framework to motivate the hedonic wage model (Jones-Lee, 1974; Rosen, 1974). However, this framework assumes risk is exogenously determined and independent of individual actions, an assumption that is not likely to hold true in the context of many types of workplace risks. Interpretation of the hedonic price for risk as a willingness to pay (WTP) measure now requires a set of assumptions about how self-protection affects both ex-ante probabilities and ex-post outcomes (Shogren and Crocker, 1991, 1999; Crocker and Shogren, 2003). In addition, the failure to account for worker's ability to self-protect in the hedonic wage model result in upwardly biased estimates of the true wage/risk tradeoffs (Shogren and Stamland, 2002). Shogren and Stamland (2006) propose a GMM framework as a partial solution, but the data needed to implement this approach is not available in our application.

Thus, conceptually, VSL estimates in this research, and all existing studies employing a hedonic wage approach to estimate the VSL, present an upper-bound on WTP estimates under the conditions as described in Shogren and Crocker (1991) and Shogren and Stamland (2002). This of course assumes no other systematic biases are introduced through the empirical implementation. Yet there are many empirical challenges when using hedonic wage models to identify risk premiums, and some important ones parallel those of hedonic property value models attempting to isolate the value of spatially-related amenities. In the twenty years since Atkinson and Crocker's (1987) explicit admonishment of "data mining (specification searching) to obtain desired signs as well as the selective reporting of unrepresentative results" (p. 28, 1987), the literature estimating the value of spatially-explicit environmental amenities through housing markets has struggled with identification strategies in the presence of unobserved spatial characteristics. The parallel in the hedonic wage literature comes directly from difficulty in obtaining data on job characteristics, and some fragility in our estimates reflect this difficulty. However, our study is certainly not unique in these regards and the issue likely extends to previously published hedonic wage studies in the VSL literature relying on similar data and methods.

More recently, the use of quasi-experimental estimation strategies have been employed in the housing literature (e.g., Chay and Greenstone, 2005; Petrie and Taylor, 2006; Pope, 2008). Unfortunately, the availability of such identification strategies are not immediately apparent for our application. Thus, we directly report the sensitivity of our results to tightly parameterized models which include fixed-effects for 63 occupations and 233 industries, and simultaneously allow for error correlation among all workers assigned to the same risk-group.[3]

2. Heterogeneity of workplace risks and compensating wage differentials

The US Bureau of Labor Statistics (BLS) began an annual Census of Fatal Occupational Injuries (CFOI) in 1992 that substantially improved on previous data collection efforts.[4] The CFOI is a complete

[3] Kuminoff et al. (2009) conduct a simulation analysis and find that adding spatial fixed effects for house location substantially reduces the bias from omitted spatially-related variables in cross-sectional housing data. This directly parallels our use of disaggregated occupation/industry fixed-effects.

[4] Prior to the CFOI, workplace mortality data varied greatly across States in both completeness and quality of the information reported (see Drudi, 1997).

census, verified by at least two independent documents such as death certificates and worker's compensation reports to assure that the incident meets the criteria for inclusion as a work-related incident. To be considered a work-related incident, the worker had to be involved in normal work activities, carrying out usual duties, or simply be in work status (BLS, 1998). In the period of our study (1992–1997) there were over 37,000 deaths. The CFOI categorizes each of these deaths according to the event leading to death using the Occupational Injury and Illness Classification System (OIICS, 1992). The CFOI also includes a brief narrative description providing details on the incident. While these descriptions vary greatly in content and detail, they enable further classification of deaths in many cases.

Workplace deaths may be broadly categorized as arising from three types of events: traditional workplace accidents, transportation-related accidents, and violent assaults. We label events such as contact with objects or equipment, falls, electrocution, or fires as "traditional" accidental deaths since these are the types of deaths most commonly thought of when considering policy initiatives to improve workplace safety. Deaths are classified as transportation-related when the worker was a driver or passenger in a motor vehicle accident or was hit by a motor vehicle of any type. Motor vehicles include, for example, trucks, trains, airplanes, boats, tractors, and forklifts. Transportation deaths occurring while an employee is commuting to work from home, and vice-versa, are not counted as workplace deaths by CFOI. However, transportation deaths occurring while "at work" (e.g., a delivery driver en route making deliveries) or occurring while traveling "for work," such as traveling to a conference, are included in the data.

The last major category of workplace fatalities arises from violent assaults. Because the nature of these violent acts varies, the CFOI coding and narrative details were used by Scotton (2000) to further classify violent acts. First, deaths that were self-inflicted (suicides), or the result of personal, non-work-related conflicts or revenge were identified.[5] The remaining violent acts fall into five broad categories: deaths relating to robbery; deaths relating to carrying out normal work-related duties such as deaths of police, detectives, guards, and workers murdered during disputes with customers, clients, patients, or tenants; deaths arising from violence between co-workers arising from incidents such as disputes over pay; deaths from terrorism-type incidents (e.g., a bombing where the narrative indicated the target was society at large and not the particular person who died); and finally, deaths from an unknown motive. This last category, "unknown motive", includes remaining acts of violence that could not be put into any of the above categories perhaps reflecting random violence or just a lack of detail in the CFOI narrative.[6]

Table 1 reports the average annual deaths by type of event from 1992 to 1997. As Table 1 indicates, traditional accidents account for 39.5% of all workplace deaths in the US during our study period while transportation-related deaths account for 41% of total deaths. Overall, violent assaults (homicides) account for approximately 20% of average annual workplace deaths during our study period. Note, the majority of homicides are attributable to robbery or had an unknown motive, while only 13% of homicides arise in the "line of duty" such as would be the case for police officers.

The inherent heterogeneity in workplace deaths has not yet been recognized in the empirical hedonic wage literature. Labor market studies simply estimate hedonic wage functions in which wages are assumed to be a function of a single measure of the risk of death from all sources. Motivating this approach is the framework suggested by Jones-Lee (1974) (see also Rosen, 1974) which assumes a worker's utility is defined over wages, w, in either a healthy state, $U^H(w)$ or an injured state $U^I(w)$. A von Neumann–Morgenstern expected utility is then defined over wealth and the probability of an injury, ρ as follows:

$$E(U) = (1-\rho)U^H(w) + \rho U^I(w). \tag{1}$$

[5] For example, any CFOI narratives identifying the assailant as an (ex-) spouse, (former) friend or an immediate relative, or indicating the worker was killed while committing a violent act or resisting arrest were placed in the category "personal conflict."

[6] Examples from the 'unknown' category are a cashier from an all-night gas station whose body is found miles away or a driver found dead near his abandoned truck, "possibly a result of a bad drug deal." These may be acts of robbery or personal revenge, but it is unknown for sure. They certainly indicate the general, perhaps random, violence in US society.

Table 1
Average annual deaths by type, 1992–1997.

	Number of deaths (percent of total)	Percent of subgroup
Traditional	2493 (39.5)	
Contact w/objects and equipment	1033	41.4
Falls	657	26.3
Bodily reaction/exertion	21	0.8
Exposures	588	23.6
Fires and explosions	194	7.8
Transportation	2589 (41.0)	
Motorized vehicle or equipment	2068	79.9
Railway or water craft	186	7.2
Aircraft	322	12.4
Mode not specified	13	0.5
Assaults and violent acts	1218 (19.3)	
Self-inflicted (suicide)	215	17.7
Personal conflict	49	4.0
Robbery	338	27.7
Line of duty	163	13.4
Co-worker conflicts	56	4.6
Terrorism	24	2.0
Unknown motive	373	30.6
Other events/exposures	13 (0.2)	
Total deaths	6313	

Source: Computed by authors using the CFOI.

If the probability of injury, ρ, is a probability of fatal injury, U^I is considered utility in the state of death, and may be thought of as representing the current period's utility over wealth bequests to heirs. In the standard formulation, utility is unaffected by the circumstances surrounding death and there is only one risk, ρ. However, expected utility in (1) is easily extended to allow utility to be distinguished by the circumstances surrounding death:

$$E(U) = \rho_N U^H(w) + \sum_{i=1}^{N-1} \rho_i U^{I_i}(w), \tag{2}$$

where ρ_N is the probability of being in the healthy state, and there are $N-1$ mutually exclusive probabilities of being in an injured state caused by event i, ρ_i. We assume $\rho_i < 1$ for all i, $\Sigma_{i \neq N}\, \rho_i < 1$, $\rho_N = 1 - \Sigma_{i \neq N}\, \rho_i$, U^H is defined as before, and U^{Ii} is the current period utility over wealth bequests to heirs and this utility is indexed by i so that utility is affected by the type of risk faced. Assuming injury risks are mutually exclusive, and that increases in ρ_i decrease ρ_N by an equal amount, the willingness to accept compensation (i.e., additional wages), v, for an additional unit of risk is defined by[7]:

$$(\rho_N^0 - \delta_1) U^H(w+v) + (\rho_1^0 + \delta_1) U^{I_i}(w+v) + \sum_{i=2}^{N-1} \rho_i^0 U^{I_i}(w+v) = \rho_N^0 U^H(w) + \sum_{i=1}^{N-1} \rho_i^0 U^{I_i}(w), \tag{3}$$

where a superscript 0 indicates an initial level of risk and δ_1 is a positive increment to ρ_1. The marginal willingness to accept compensation for a change in ρ_1 is then given by:

$$\frac{\partial v}{\partial \delta_1} = \frac{U^H(w+v) - U^{I_1}(w+v)}{(\rho_N - \delta_1)\frac{\partial U^H}{\partial v} + (\rho_1 + \delta_1)\frac{\partial U^{I_1}}{\partial v}} > 0 \tag{4}$$

[7] We assume injury risks are mutually exclusive. However, it could be the case that there are complementarities in risk reductions and that actions taken to reduce ρ_i also reduce ρ_j.

and is convex in risk (i.e., $\partial^2 v/\partial \rho^2 > 0$).[8] A marginal willingness to pay function (i.e., the willingness to receive lower wages) for risk reductions is defined analogously.

As made clear by (4), differences in the marginal willingness to accept compensation for a change in any single risk is driven by utility differences among the heterogeneous risks. With M heterogeneous firms and N heterogeneous workers in a competitive labor market, the equilibrium hedonic wage function will then be differentiated by risk: $W(\rho_1, \ldots, \rho_{N-1})$.

While conceptually simple, defining differentiated risks in an empirical context is not as easily done. In the context of workplace deaths, evidence suggests that workplace violent assault risks might affect utility differently than traditional or transportation-related fatality risks. Deaths from traditional events are, for the most part, readily understood to be part of dangerous jobs such as mining. While the death itself is an anomaly, the events leading to it are often routine aspects of the work environment, with little perceived dread and a high degree of perceived controllability. The same can be said for transportation-related accidents as indicated by Slovic et al. (1980). On the other hand, violent assaults leading to death are judged by individuals to have a great degree of dread while offering little controllability (Slovic et al., 1980). Jones-Lee and Loomes (1995), Sunstein (1997), and Hammitt and Liu (2004) all report increased willingness to pay to avoid risks that are characterized by dread or that are not controllable. As such, we expect that homicides might be compensated differently in the workplace. We have no strong priors regarding the differentiation among transportation and traditional sources of risk and so we empirically test all three types of risks for aggregation properties. In the next section, we describe construction of the risk measures and the labor force sample used to test our hypotheses.

3. Data

For each workplace death in the US, the CFOI not only records the circumstances surrounding death, but also each deceased worker's 3-digit occupation code based on the 1990 Census Occupation Classification System and the 4-digit industry code from the Standard Industrial Classification (SIC) Manual, 1987 edition. We aggregate the occupations into 22 categories and the industries into 23 categories. These 506 occupation/industry pairs account for every possible occupation in every possible industry in the US as listed by the BLS. To create risk rates, we divide the average number of deaths recorded for each occupation/industry pair during the 1992–1997 period by the average number of workers in each occupation/industry pair during the period 1991–1996. This latter data is provided by the BLS Office of Occupational Employment Statistics, Industry-Occupational Employment Matrix. The advantage of using annual averages, rather than a single year, is to avoid focusing on unusual circumstances such as a large industrial accident, and to smooth out anomalies in the employment survey data. To avoid small-sample problems, occupation/industry pairs with fewer than 500 workers nationwide were excluded, leaving 492 occupation/industry risk rates available for analysis.[9] Ten percent of the occupation/industry pairs had no deaths occur during the 6-year study period.

To create a consistent measure of the risk of death for inclusion in the hedonic wage model, only deaths that are likely to be considered as part of a wage-negotiation between an employer and employee are included. As such, violent assault deaths that were self-inflicted (suicide) or due to a personal conflict are removed because they could not be construed as systematically related to work.[10] Deaths which took place more than 30 days after the workplace incident were also excluded from the data set.[11] Furthermore, deaths of those in the military and the self-employed were removed. Self-employed workers, while representing less than 10% of the US labor force, account for about 20%

[8] This requires the assumption that the utility function is more sensitive to changes in wealth when in the healthy state as compared to an injured state, i.e., that $\partial U^H/\partial w > \partial U^I/\partial w$ and $|\partial^2 U^H/\partial w^2| > |\partial^2 U^I/\partial w^2|$.

[9] The mean number of workers per occupation/industry pair was 234,000 with a median of 59,000. Forty-seven occupation/industry pairs (less than 10%) had fewer than 5000 workers, while 27 pairs (5%) had over 1 million workers.

[10] We could find no evidence that these types of deaths occur more frequently in any specific occupation or industry, indicating there is no systematic bias introduced by removing these deaths from our database.

[11] While the vast majority of deaths recorded in the CFOI occurred within 30 days of the injury (over 97%), some deaths occurred as many as 25 or more years after the injury.

of the fatal injuries at work (Personick and Windau, 1995). The self-employed often face a different work environment than do wage and salary workers and are generally excluded from workers' compensation coverage (Drudi, 1997).[12] Including self-employed workers in the development of the risk measures would increase risk rates for some industry-occupation pairs while decreasing the rate for others and thus bias estimates of the wage-risk relationship for the non-self-employed workforce.

After removing each of the types of deaths discussed above, there were approximately 4500 deaths annually that would plausibly be incorporated into wage negotiations between workers and firms. Using this annual average number of deaths, a risk rate undifferentiated by the type of death, r^u, is constructed for each occupation, o, within each industry, i, as follows:

$$r^u_{oi} = \frac{D_{oi}}{W_{oi}}, \tag{5}$$

where D_{oi} is the total average annual number of deaths occurring in a specific occupation within a specific industry, and W_{oi} is the total number of workers in the same occupation/industry pair.[13] For ease of exposition, the subscripts oi will be dropped as it is understood that a risk-rate is created for each of 22 occupations within each of 23 industries as described earlier.

Risk rates are also created that is differentiated by the three major categories of deaths: traditional accidental deaths, r^{td}, transportation-related accidents, r^{tr}, and violent assaults, r^{v}. Each of these risk rates is created for each of the occupation/industry pairs. The risks are computed in the same manner as r^u, except the numerator simply includes only the average annual deaths for each type of risk. Thus, $r^u = r^{td} + r^{tr} + r^{v}$.

3.1. Labor market sample

Hedonic wage models are estimated by matching the risk measures to a sample of 43,261 non-self-employed, full-time, civilian workers from the 1996, 1997, and 1998 Current Population Survey (CPS), a commonly used data source in hedonic wage studies (Viscusi and Aldy, 2003). Our sample is drawn from individuals interviewed in their fourth month in-sample (referred to as the out-going rotation group) to avoid duplicate observations for the same individual and only includes individuals who reported their earnings.[14] The sample is limited to workers earning $630 or more per week, which is considered the high-wage segment of the labor market (Bernstein and Hartmann, 2000).[15] We limit the sample to the high-wage segment for several reasons. First, the lowest-paying entry-level jobs are often those that entail very high violent risks (above the 95th percentile), e.g., cashiers/salespersons, stock handlers and baggers, service station workers, non-construction laborers, and bus/taxi drivers.[16] These low-paying entry-level jobs are also generally occupied by individuals with more limited labor-force opportunities and have larger proportions of immigrants to the US (Bucci and Tenorio, 1997). While information on immigrant population by detailed occupation is difficult to obtain, anecdotal evidence suggests that this population may be over-represented in occupations with high violent risk. For instance, a recent survey indicated ninety percent of the taxi cab drivers in New York City are immigrants (Holt and Paradise, 2002). General impressions are that immigrant workers have less opportunity for advancement and to move into higher-paying, higher-skilled jobs

[12] For example, the self-employed are more likely to be victims of homicide, perhaps because of those in retail business (e.g., small retail shop owners) working alone during evening hours, while self-employed workers in the construction trades are less likely to be involved in a fatal accident because they tend to be concentrated in small construction sites (Personick and Windau, 1995, p. 58). In addition, over 40% of deaths for the self-employed are farm-related accidents.

[13] To create consistent risk rates, we use only data from the Industry-Occupational Employment Matrix on the average annual level of employment for all civilian, non-self employed (or family) workers in each occupation/industry pair.

[14] Approximately 26% of the CPS sample we drew did not report wages. Although the BLS imputes wages for workers who refuse to report them, the methodology used by the BLS to impute wages may introduce bias in the coefficient estimates (see Bollinger and Hirsch, 2006).

[15] The low-wage labor market segment is defined as those earning at the poverty level for a family of four in 1997 ($16,500 or approximately $300 per week), and the medium-wage sector is defined as those earning up to twice the poverty level, which is $630 per week ($31,500 per year).

[16] In fact, taxi drivers have the highest violent assault risks of all occupations in the U.S.; nearly four-times that of police officers (Knestaut, 1997).

as native counterparts.[17] During the period 1990–1998, median immigrant earning were $13,000 (Camerota, 1999) and Bucci and Tenorio (1997) find large immigrant-native wage differentials. Unfortunately, we do not have available information on characteristics such as English proficiency and immigrant status. Limiting our sample may help alleviate the bias introduced by these unobservable characteristics.

Related to the above, if labor market segmentation based on skill levels exists, with more job mobility among jobs within a segment than across segments, models of wage determination should not specify the same data generating process across segments (Bernstein and Hartmann, 2000; Harrison and Sum, 1979). This type of segmentation makes it problematic to identify a single wage/risk premium for a general sample of workers, especially when the risk of interest occurs primarily in occupations characterized by low skills/low wages. This problem is exacerbated if workers in the low-skill/low-wage segment have little ability to adjust to changes in wages and working conditions as the equilibrium conditions require.

Sample characteristics are reported in Table 2. The mean (median) gross weekly wage for our sample is $996 ($874). Overall, the mean risk of death for the sample of workers (using r^u) is approximately 0.5×10^{-4}, the median is 0.2×10^{-4}, and the range is 0 to 35.5×10^{-4}. There are 136 observations (0.3%) in our sample which are assigned a zero risk rate for r^u and 4525 observations (10%) with a risk rate greater than 1.0×10^{-4} for r^u. The 90th and 95th percentile risk for our sample are 1.04×10^{-4} and 2.45×10^{-4}, respectively. Table 2 also reports the range of the risks and the percent of total risk (r^u) that each of the differentiated risks represents for each of the four risk measures for our sample of workers. The pattern of deaths for our sample of workers generally follows that reported in Table 1 with transportation and traditional accidents comprising the majority of deaths, although within these two risks, transportation is a larger share of overall risk as compared to the entire workforce. For comparison, we also draw a broader sample from the CPS of approximately 116,000, using the same out-going rotation groups as our high-wage sample but including all full-time workers earning at least minimum-wage. The mean, median and range of risks of risks faced by our sample of high-wage workers is the same as that for the broad CPS sample.[18]

The workforce sample somewhat under-represents women and minority populations, while also having a higher level of education and percentage married as compared to the general population. As compared to the broad CPS sample of all workers earning at least minimum-wage, the high-wage sample has a higher proportion of workers who hold an undergraduate degree, are white, male, and married. In addition, the high-wage sample has a higher proportion of workers who are salaried, work overtime, are in a union, and live in a large metropolitan area. High-wage workers are distributed across broad industry categories in a similar fashion as the larger sample of all workers, but are more concentrated in professional occupations and less concentrated in service and laborer occupations.

4. Empirical model and results

The basic model used to investigate the wage/risk relationship for our sample of workers, assuming initially that risks are undifferentiated, can be written as:

$$\ln(wage_k) = \alpha + \beta_u r_k^u + \beta_{u2}(r_k^u)^2 + \sum_{n=1}^{N} \delta_n X_{kn} + \sum_{m=1}^{M} \lambda_m D_{km} + \varepsilon_k, \tag{6}$$

where r^u is the undifferentiated risk of death for the kth worker, X represents the N human capital, demographic and job characteristics assumed to influence wages ($N=19$ in our model), D represents the dummy variables identifying the industry and occupation of the worker, α, β, δ, and λ are coefficients to be estimated, and ε is the error term. All models reported compute robust standard

[17] US Commission on Civil Rights, Presentation on Civil Rights Issues Facing Immigrant Communities, December 13, 2002, http://www.usccr.gov/pubs/tragedy/imm1202.htm.

[18] The sample of high wage earners represents 439 of the possible 492 occupation/industry pairs upon which our risk measures are constructed, while the sample of 116,000 workers represents 469 of the 492 risk groups.

Table 2
Variable definitions and sample characteristics (N=43,261).

Var. name	Definition	Mean	Std. dev.
Wage and risk characteristics			
WAGE	Gross weekly wage	995.78	378.47
r^u	Total deaths per 10,000 workers. Range of risk in brackets	0.4895 [0–35.54]	1.0112
r^{td}	Deaths per 10,000 workers from traditional sources. Range of risk in brackets. Percent of total risk that r^{td} represents, on average, for sample is in parentheses[a]	0.2031 [0–17.77] (29.5)	0.5675
r^{tr}	Deaths per 10,000 workers from transportation sources. Range of risk in brackets. Percent of total risk that r^{tr} represents, on average, for sample is in parentheses[a]	0.2354 [0–14.21] (51.6)	0.5912
r^{v}	Deaths per 10,000 workers from all violent assaults. Range of risk in brackets. Percent of total risk that r^{v} represents, on average, for sample is in parentheses[a]	0.0510 [0–4.80] (18.9)	0.1405
NONFATAL	Nonfatal workplace injuries by industry per 100 workers	7.16	4.34
Worker characteristics			
AGE	In years	41.7	9.5
UGDEG	1 if individual has a four-year college degree	0.50	0.50
COLLEGE	1 if individual attended college	0.26	0.44
HSGRAD	1 if individual graduated from high school (or earned GED)	0.21	0.41
WHITE	1 if individual is white	0.86	0.35
HISPANIC	1 if individual is of Hispanic origin	0.04	0.20
BLACKNH	1 if individual is black, non-Hispanic	0.05	0.23
OTHRACE	1 if individual is not black, white or Hispanic	0.04	0.21
FEMALE	1 if individual is a female	0.33	0.47
MARRIED	1 if individual is married	0.72	0.45
Job characteristics			
SALARY	1 if individual works for a salary	0.67	0.47
WORKOT	1 if individual usually works >40 h/week	0.53	0.50
UNION	1 if individual is a union member	0.26	0.44
MW	1 if individual lives in Midwest region	0.26	0.44
SOUTH	1 if individual lives in southern region	0.26	0.44
WEST	1 if individual lives in western region	0.25	0.44
MIDSIZE	1 if individual lives in MSA with pop. >25,000 and <1 million	0.18	0.39
LRGMSA	1 if individual lives in MSA with population > 1,000,000	0.58	0.49
URBAN	1 if individual lives in the central city of MSA	0.22	0.42
CONSTIND	1 if individual is in the construction industry	0.06	0.24
MANUFIND	1 if individual is in the manufacturing industry	0.21	0.41
AGIND	1 if individual is in agriculture, forestry, fishing, or mining industry	0.02	0.14
TCUIND	1 if worker is in transportation, communications or utilities industry	0.10	0.30
TRDIND	1 if individual is in a wholesale or retail trades industry	0.10	0.30
SERVIND	1 if individual is in a service industry (including finance, insurance and real estate)	0.41	0.49
PUBIND	1 if individual is in the public sector	0.09	0.29
PROFOCC	1 if individual is in a professional occupation	0.54	0.50
CRAFTOCC	1 if individual is in a craftpersons occupation	0.13	0.34
TECHOCC	1 if individual is in a technical occupation	0.21	0.41
SERVOCC	1 if individual is in a service occupation	0.03	0.18
LABOROCC	1 if individual is operator or laborer	0.08	0.28
FARMOCC	1 if individual is in a farming, forestry or fishing occupation	0.003	0.06

[a] To compute the percent of total risk that each risk type represents, we compute the percent that each risk contributes to total risk (r^u) for each worker in the sample, and then report the mean of that percent calculation.

Table 3
Select results, hedonic wage models with undifferentiated risks (dependent variable is natural log of gross weekly wages, $N=43{,}261$).[a]

	Coefficient (standard error)		
	Model 1	Model 2	Model 3
r^u	0.0282*** (0.0093)	0.0320*** (0.0083)	0.0186*** (0.0057)
$(r^u)^2$	−0.0009*** (0.0003)	−0.0009*** (0.0002)	−0.0005*** (0.0002)
Occupation/industry category controls in the model	6/7	22/23	63/233
Model R^2	0.249	0.277	0.310
VSL, in millions (std. error)	$8.7 (2.88)	$9.9 (2.58)	$5.8 (1.78)

[a] Robust standard errors are reported that allow for correlation among observations assigned the same risk rate. Models contain all covariates listed in Table 2, varying only by the number of industry and occupation control variables as indicated in the table. VSL point estimates are computed assuming a wage level of $630 and the sample mean risk, $r^u=0.4895$.
*** Coefficient estimate is significant at the 1% level.

errors that allow for correlation among observations which are assigned the same value for r^u (i.e., all observations in the same occupation/industry pair).

Each of the 19 worker and job characteristics included in the models are described in Table 2. Age, educational attainment, race, gender and marital status are included in the models, as are controls for job characteristics such as whether or not workers earn a salary, usually work more than 40 h per week, are union members or covered by a union contract. Also included in the models are controls for the region of the US in which the worker resides, as well as categorical variables indicating the size of the Metropolitan Statistical Area (MSA) in which the workers lives. Non-fatal risk of injury on the job is also included in the control variables, but unfortunately this measure is only available at an industry-level and not available for workers in the public sector.[19] As a result, public safety workers such as firefighters and police are excluded from our sample of workers.

The models also control for the occupation and industry of the worker. Most hedonic wage models rely on risk measures where the primary variation arises from only inter-industry risk differentials without consideration of the worker's occupation. This creates measurement error and induces multicollinearity that renders most of the past hedonic wage/risk estimates unreliable (Leigh, 1995; Mrozek and Taylor, 2002; Black and Kniesner, 2003). We test the robustness of our models to various levels of industry and occupation controls. In the base models, we include dummy variables indicating the six major occupation and seven major industry groups as listed in Table 2. We also report models with dummy variables for each occupation and each industry category used to create the risk measure (22 occupations and 23 industries) and a final model that includes controls for a disaggregated set of 63 occupations and 233 industries. Furthermore, in all models, recall that we cluster our standard errors to allow for correlation among all observations within a specific industry/occupation pair upon which our risk measure is created. As such, our models tightly control for general occupation and industry wage differentials relative to the risk measure construction method, thus giving more confidence that our resulting risk-related coefficient estimates reflect the impact of risk level on wages, and not something more general about the occupation or industry affiliation of the worker.

Table 3 reports the coefficient estimates for undifferentiated risks based on the model presented in Eq. (6). Full results for each model are available in Table A.1. In general, non-risk coefficient estimates are statistically significant, of the expected signs, and are stable across all models we estimate.[20] As indicated in Table 3, there is a positive, significant relationship between undifferentiated fatal job risks

[19] The non-fatal workplace accident risk was created from published reports of total nonfatal workplace accidents for private industry in 1997 as published by the BLS (http://www.bls.gov/iif/oshwc/osh/os/ostb0642.pdf). We computed the rate at 2-digit SIC level. The correlation coefficient between the non-fatal risk measure and undifferentiated fatal risks, r^u, is 0.16.
[20] The exception is the coefficient estimate for non-fatal risk, which is never significantly different from zero for our sample of high-wage earners.

and wages. The fatal-risk coefficients are robust to the number of industry and occupation control variables, and although not reported here for succinctness, the models are robust to excluding the highest risk workers (95th percentile or higher) and those with zero fatal risks. Point estimates for the VSL, based on the mean risk for the sample are \$6–10 million. This range encompasses the VSL point estimate of \$8.4 million reported by Viscusi (2004, p. 32, footnote 25) who uses similar occupational-by-industry risk data, but reports models only with risk entered linearly and uses only occupation control variables. Note however, our point estimates fall to between \$3.4 and \$6.0 million (and are significantly different than zero) if we modify our base model to more closely match Viscusi's (2004) modeling choices and include only a linear risk term and only occupational control variables.

To examine whether aggregation of risks is a mis-specification, we estimate models similar to that presented in Eq. (6), but which include risk disaggregated by type. We then test for linear aggregation of the risk coefficients. *F*-Statistics are computed with q numerator degrees of freedom and m denominator degrees of freedom, where q equals the number of restrictions, and m equals the number of independent sampling units. Since we allow for correlation among observations arising from the same occupation/industry group (a "cluster"), m is equal to the number of clusters. For example, if we wish to test the hypothesis that all risks are additive, $r^u = (r^{td} + r^{tr} + r^{\nu})$, implying the following model:

$$\ln(wage_k) = \alpha + \beta_u(r_k^{td} + r_k^{tr} + r_k^{\nu}) + \beta_{u2}(r_k^{td} + r_k^{tr} + r_k^{\nu})^2 + \sum_{n=1}^{N}\delta_n X_{kn} + \sum_{m=1}^{M}\lambda_m D_{km} + \varepsilon_k, \tag{7}$$

we estimate the following unrestricted model:

$$\ln(wage_k) = \alpha + \beta_{td} r_k^{td} + \beta_{td2}(r_k^{td})^2 + \beta_{tr} r_k^{tr} + \beta_{tr}(r_k^{tr})^2 + \beta_{\nu} r_k^{\nu} + \beta_{\nu 2}(r_k^{\nu})^2 + \beta_{\text{int}} 2(r_k^{td} r_k^{tr} + r_k^{td} r_k^{\nu} + r_k^{tr} r_k^{\nu}) + \sum_{n=1}^{N}\delta_n X_{kn} + \sum_{m=1}^{M}\lambda_m D_{km} + \varepsilon_k, \tag{8}$$

and test the following restrictions:

$$H_0^{T5} : \beta_{td} = \beta_{tr} = \beta_{\nu} \ \text{and} \ \beta_{td2} = \beta_{tr2} = \beta_{\nu 2} = \beta_{\text{int}} \tag{9}$$

against the alternative that at least one of the above is not equal. Note, the quadratic interaction coefficient, β_{int}, included in (8) arises from the expansion of the term $(r^{td} + r^{tr} + r^{\nu})^2$ in Eq. (7). For the three models presented in Table 3, the following aggregation combinations are tested separately: $\{r^{td} + r^{tr} + r^{\nu}\}$, $\{r^{td} + r^{tr}\}$, $\{r^{td} + r^{\nu}\}$, and $\{r^{tr} + r^{\nu}\}$. When we test the aggregation of only two of the three risks, the risk not aggregated is also included in the model as a separate control variable. For example, when testing whether r^{td} and r^{ν} can be aggregated, r^{tr} is included as a control variable in the model. Results across models generally indicate that we can reject aggregation of violent risks with traditional risks, transportation risks or the sum of traditional and transportation risks at the 5% level or better, but cannot reject aggregation of transportation and traditional risks. The exception is the model which utilizes 22 occupation and 23 industry control variables. In this model, we can only reject aggregation of violent and transportation risks at the 5% level, and aggregation of transportation and traditional risks are rejected. Overall, the aggregation tests imply the following hedonic wage equation:

$$\ln(wage_k) = \alpha + \beta_t r_k^t + \beta_{t2}(r_k^t)^2 + \beta_{\nu} r_k^{\nu} + \beta_{\nu 2}(r_k^{\nu})^2 + \sum_{n=1}^{N}\delta_n X_{kn} + \sum_{m=1}^{M}\lambda_m D_{km} + \varepsilon_k, \tag{10}$$

where $r^t = (r^{td} + r^{tr})$ and all other variables are previously defined.

The first three columns in Table 4 present results related to the risk coefficients for the model in Eq. (10). Again, non-risk related variables are stable and consistent across models. Note, in Table 4 and the remainder of this paper, we no longer refer to a VSL since we are individuating risk reduction values by the type of risk faced. We instead use the more precise phrase "value of a risk reduction" (VRR) and superscript VRR estimates with the type of risk being considered (r^t or r^{ν}). As indicated in the first three columns of Table 4, aggregated traditional and transportation risks, r^t, are precisely estimated and of the expected sign, regardless of the number of industry and occupation control variables included in the model. In addition, the point estimates for the VRR^t are comparable to the

Table 4
Hedonic wage regressions with differentiated risks (dependent variable is natural log of gross weekly wages)[a].

Risk measure	All high-wage workers			Less administrative and sales		
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
$r^t = (r^{tr} + r^{td})$	0.0222** (0.0102)	0.0346*** (0.0090)	0.0200*** (0.0059)	0.0177* (0.0100)	0.0262*** (0.0098)	0.0079* (0.0046)
$(r^t)^2$	−0.0006* (0.0003)	−0.0011*** (0.0003)	−0.0006*** (0.0002)	−0.0005 (0.0003)	−0.0008** (0.0003)	−0.0002 (0.0002)
r^v	0.2424*** (0.0881)	0.0141 (0.0501)	0.0241 (0.0535)	0.1117* (0.0590)	0.0957** (0.0445)	0.0949** (0.0404)
$(r^v)^2$	−0.0793*** (0.0288)	−0.0057 (0.0175)	−0.0144 (0.0147)	−0.0376* (0.0192)	−0.0342** (0.0135)	−0.0350*** (0.0104)
Number obs.	43,261	43,261	43,186	30,882	30,882	30,882
Occupation/industry category controls	6/7	22/23	63/233	6/7	22/23	63/233
R^2	0.250	0.277	0.310	0.220	0.238	0.288
VRR^t [1.0] (std. error)	6.8 m (3.12)	10.6 m (2.76)	6.1 m (1.81)	5.4 m (3.04)	8.0 m (2.97)	2.4 m (1.39)
VRR^v [1.0] (std. error)	73.8 m (26.8)	4.5 m (15.3)	7.1 m (16.4)	34.2 m (18.1)	29.3 m (13.7)	29.0 m (12.5)
$VRR^t = VRR^v$: prob > F[b]	0.0144	0.6825	0.9595	0.1272	0.1519	0.0390

[a] Robust standard errors are reported that allow for correlation among observations assigned the same risk rate. Models contain all covariates as listed in Table 2. VRR point estimates are reported in millions and are computed assuming an average wage level of $630 and the mean level of risk (r^t or r^v) for the sample.
[b] Tests reported are computed assuming a common risk level for both r^v and r^t (equal to the mean violent risk for the sample). Results are not qualitatively sensitive to the choice of risk-level.
* Coefficient estimates that are significant at the 10% level.
** Coefficient estimates that are significant at the 5% level.
*** Coefficient estimates that are significant at the 1% level.

point estimates reported in Table 3 based on undifferentiated risks.[21] The coefficient estimates for violent risks are all of the expected sign, however, they are not robust to models including greater specificity in occupation and industry controls. The first model in Table 4, which includes only six occupation and seven industry controls indicates a statistically significant relationship between violent risks and wages, and result in a very large point-estimate for the value of a one-unit risk reduction, $74 million. This point estimate of the VRR^v is significantly different than the point estimate of VRR^t at the 2% level.[22] However, when additional occupation and industry control variables are included in the model, the estimated coefficients for violent risk are reduced in magnitude and are no longer statistically significant (Table 4, models 2 and 3).

To explore these results, we estimate models with the six occupation dummy variables listed in Table 2 while including either 23 or 233 dummy variables for the worker's industry. The estimated coefficient for traditional and violent risks are remarkably stable in magnitude and significance as compared to model 1 in Table 4. However, if instead models are estimated with the seven industry dummy variables listed in Table 2 and either 22 or 63 occupation dummy variables, the violent risk measure falls in magnitude and is no longer statistically significant. It appears that unobserved occupation characteristics are correlated with violent risks faced on the job, regardless of the industry in which a worker works.[23] We thus examine occupation groups for influential categories on the violent risk coefficient estimates. Using dfbeta influence statistics, and a cutoff value of $2/\sqrt{n}$ (Belsley et al., 1980), two broad occupation categories account for nearly 40% of the most influential observations for the violent risk coefficient estimates. These occupations categories are "Executive and Administrative Positions," and "Marketing and Sales Occupations." These two occupation categories are routinely in contact with the public, and thus more susceptible to violent encounters. In addition, there may be potentially important unobservable occupation characteristics that are correlated with violent risk, such as the hours of business operations. For instance, sales jobs which involve working at night would most likely be at higher risk of violent assault than those working during regular day-light business hours. However, we only have information on the hours usually worked, not when those hours occur. Our detailed occupation dummy variables control for some of this variation by distinguishing, for instance, between "finance and business service sales representatives" who are unlikely to work in the evenings and "retail and personal services sales representatives" which are more likely to work during a variety of day and evening hours.

The last three columns of Table 4 report models in which we exclude workers in these two occupation categories. Traditional risks are less precisely estimated in these models, although the point-estimates for VRR^t are still significantly different from zero at the 10% level in all models. Violent risks are significantly different from zero in all three models, and result in point-estimates of the VRR^v evaluated for a one unit change in risk of approximately $30 million.[24] Point estimates of VRR^v are significantly larger than the VRR^t at the 10% level in all three models in a one-tailed test, and significantly different in a two-tailed test for the last model reported in Table 4 which has the most detailed occupation and industry control variables.[25]

[21] These results suggest that including violent assault risks in an undifferentiated total measure of fatality risk has not likely been a source of significant bias in previous hedonic wage studies. Estimated wage premia for traditional sources of workplace risk, net of homicide risks, are consistent with previous estimates for undifferentiated risks (e.g., Viscusi and Aldy, 2003; Viscusi, 2004) and are generally stable to the level of detail in the occupational and industry control variables.

[22] Tests are performed on point estimates of the VRR^t and VRR^v computed at a wage of $640 and either the mean traditional risk or the mean violent risk for the sample.

[23] This point is further highlighted by models estimated using the large CPS sample of all workers earning at least minimum wage ($N = 115{,}881$). In these models, violent risk coefficients have the opposite signs, indicating wages decrease with increased violent risks, and these results are statistically significant in some models.

[24] A one-unit increase in r^v from the mean risk of violent death moves a worker to the 99th percentile of violent risk. If instead, we consider a more meaningful change in risk for the sample: a one-standard deviation change in risk ($\sigma = 0.16$), point estimates of the VRR^v are approximately $5 million. From a policy analysis perspective, estimated total benefits of a risk change would be the same regardless of whether the intermediate reporting focused on the traditional one-unit change in risk (VSL) or on a VRR scaled to more appropriate risk changes.

[25] The exception is the model including 22 occupation controls and 23 industry controls. The point estimates of the VRR^t and VRR^v are not significantly different in a one-tailed test if evaluated at the mean traditional risk, but are significantly different if evaluated at the mean violent risk.

5. Concluding comments

The use of labor-markets to estimate the tradeoffs people are willing to make between income and the risk of injury is not without important criticisms, yet these types of studies are routinely used as the primary input in developing VSL point-estimates for major policy analyses (e.g., US EPA, 2005, 1999). While much attention and criticism has been given to the types of risk measures used in past studies (e.g., Black et al., 2002; Mrozek and Taylor, 2002; Scotton and Taylor, 2003; Shogren and Stamland, 2002), there has been little discussion in the revealed preference literature about the possibility that aggregating the risk of death across multiple causes is inappropriate. In this research, we focus on the circumstances surrounding a death as a possible source of risk-differentiation. Our empirical tests indicate that violent assault risks (i.e., homicide) is differentiated among workplace risks, and additional homicide risks command a significantly higher wage premium as compared to traditional sources of workplace risk such as falls, electrocution, or motor vehicle accidents.

In developing our hedonic wage models, a cautionary tale emerges however. Cross-sectional hedonic wage models rely on occupation and industry categorical variables to capture unobserved job characteristics. Yet workplace risks are measured by occupation and industry category as well. The resulting empirical challenges that arise from risk mis-measurement and omitted variables has long been recognized in the VSL literature (Leigh, 1995; Black et al., 2002; Mrozek and Taylor, 2002; Black and Kniesner, 2003; Ashenfelter and Greenstone, 2004), and recent attempts to employ panel models and/or improve risk measurement by creating occupation within industry risk measures will not alleviate this problem.[26] At a minimum, we thus report directly the sensitivity of our results to models which include fixed-effects for 63 occupations and 233 industries, and simultaneously allow for error correlation among all workers within the same risk-group.[27]

While our results are not unequivocal, they are reasonably suggestive that the willingness to accept higher wages in return for accepting more risk on the job is significantly larger for homicide risks than for traditional workplace risks. As such, the application of VSL estimates from labor market studies, which our results clearly indicate are driven by traditional sources of workplace risks (e.g., electrocution, falls, traffic accidents) to policy contexts involving reducing latent cancer risks, or premature mortality from acute asthmatic events, for example, is very likely inappropriate. Furthermore, explicit consideration of the heterogeneous values for heterogenous risks underscores the importance of moving the policy discussion from "a VSL" to valuation of marginal changes in fatality risks specific to the type of the risk affected by a policy. Our results are at least suggestive that even within the relatively narrow context of compensation for workplace fatality risks, the way we die matters.

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[26] Panel models recently implemented by Kneisner et al. (2006) and Kochi (2010) control for time-invariant worker characteristics. However, the coefficient estimates for risk are identified from workers who change occupation or industry category (and therefore change the level of risk faced), thus offering no additional identification strategy for risk coefficients.

[27] Kuminoff et al. (2009) conduct a simulation analysis and find that adding spatial fixed effects for house location substantially reduces the bias from omitted spatially-related variables in cross-sectional housing data. This directly parallels our use of disaggregated occupation/industry fixed-effects.

Appendix A

Table A.1 Full results for models presented in Table 3.

	Coefficient (standard error)		
	Model 1	Model 2	Model 3
r^u	0.0282*** (0.0093)	0.0320*** (0.0083)	0.0186*** (0.0057)
$(r^u)^2$	−0.0009*** (0.0003)	−0.0009*** (0.0002)	−0.0005*** (0.0002)
NONFATAL	0.0418 (0.2068)	−0.1029 (0.0829)	−0.0525 (0.0753)
AGE	0.0234*** (0.0023)	0.0224*** (0.0021)	0.0227*** (0.0020)
AGE2	−0.0002*** (0.0000)	−0.0002*** (0.0000)	−0.0002*** (0.0000)
UGDEG	0.1871*** (0.0109)	0.1860*** (0.0109)	0.1740*** (0.0105)
COLLEGE	0.0646*** (0.0079)	0.0530*** (0.0079)	0.0475*** (0.0076)
HSGRAD	0.0269*** (0.0077)	0.0209*** (0.0074)	0.0183** (0.0072)
HISPANIC	−0.0321*** (0.0082)	−0.0302*** (0.0080)	−0.0256*** (0.0071)
BLACKNH	−0.0405*** (0.0090)	−0.0380*** (0.0090)	−0.0347*** (0.0075)
OTHRACE	−0.0001 (0.0120)	0.0007 (0.0115)	0.0003 (0.0118)
FEMALE	−0.1230*** (0.0069)	−0.1088*** (0.0066)	−0.1029*** (0.0068)
SALARY	0.0569*** (0.0084)	0.0576*** (0.0074)	0.0529*** (0.0053)
WORKOT	0.1046*** (0.0061)	0.1021*** (0.0054)	0.1003*** (0.0060)
UNION	−0.0270 (0.0212)	0.0032 (0.0129)	0.0206** (0.0095)
MARRIED	0.0336*** (0.0054)	0.0331*** (0.0046)	0.0355*** (0.0042)
MIDWEST	−0.0238*** (0.0074)	−0.0237*** (0.0066)	−0.0223*** (0.0064)
SOUTH	−0.0292** (0.0113)	−0.0262*** (0.0096)	−0.0258*** (0.0088)
WEST	0.0015 (0.0112)	0.0060 (0.0102)	0.0052 (0.0099)
MIDSIZE	0.0479*** (0.0069)	0.0419*** (0.0054)	0.0394*** (0.0049)
LRGMSA	0.1170*** (0.0134)	0.1086*** (0.0107)	0.1004*** (0.0089)
URBAN	0.0079 (0.0052)	0.0078 (0.0051)	0.0028 (0.0039)
CONSTIND	−0.0016 (0.0190)	–[a]	–[b]
AGIND	0.0068 (0.0289)		
TCUIND	−0.0027 (0.0228)		
TRDIND	−0.0550*** (0.0195)		
SERVIND	−0.0343 (0.0217)		
PUBIND	−0.0465** (0.0204)		
PROFOCC	0.0850*** (0.0155)		
TECHOCC	0.0335* (0.0177)		
SERVOCC	−0.0099 (0.0274)		
LABOROCC	−0.0522*** (0.0147)		
FARMOCC	−0.1314*** (0.0192)		
YR97	0.0219*** (0.0029)	0.0220*** (0.0027)	0.0212*** (0.0027)
YR98	0.0582*** (0.0033)	0.0575*** (0.0031)	0.0574*** (0.0032)
Constant	5.9426*** (0.0580)	5.9823*** (0.0635)	5.9983*** (0.0549)
Observations	43,261	43,261	43,261
R-squared	0.249	0.277	0.310

[a] Results for dummy variables indicating which of the 22 occupation and 23 industry categories in which the worker worked are not included here for succinctness.
[b] Results for dummy variables indicating which of the 58 occupation and 230 industry categories in which the worker worked are not included here for succinctness.
* Significance at the 10% level.
** Significance at the 5% level.
*** Significance at the 1% level.

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