Unions and hazard pay for COVID-19: Evidence from the Canadian Labour Force Survey

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Abstract: In this article, we examine whether (and by how much) workers in Canada have been compensated for the ‘novel’ risks associated with COVID-19. We create a unique dataset from a system that scores occupations in the US O*NET database for COVID-19 exposure. We then combine those COVID exposure scores with Canadian occupational data contained in the Public Use Microdata File of the Labour Force Survey. This allows us to categorize Canadian occupations based on COVID-19 exposure risk. We find a long-tailed distribution of COVID-19 risk scores across occupations, with most jobs at the lower end of the risk spectrum and relatively few occupations accounting for most of the high COVID-19 exposure risk. We find that workers who are already more vulnerable in the labour market (i.e. youth, women and immigrants) are also more likely to be employed in occupations with high COVID-19 exposure risk. When we look at the relationship between high-COVID exposure risks in occupation and wages, we find negative compensating differentials both at the mean (negative 8%) and across the earnings distribution. However, when workers are covered by a union, they enjoy a sizeable hazard pay premium (11.7% on average) as compared to their non-union counterparts. Furthermore, we find that the moderating effects of unionization for workers at high risk of COVID exposure to be largest at the bottom of the earnings distribution (i.e. the 10th percentile of unionized earners receives a 12.3% risk
1 | INTRODUCTION

COVID-19 has exposed many social and work-related inequalities. Consider that even during a pandemic, most employees still had to show up in person to perform their job-related duties (Lu, 2020).\(^1\) Indeed, collectively we are more aware than ever before of the ‘essential’ workers who not only care for us but provide us with food and the other basic necessities of life. This gratitude has been reflected in the rhetoric of our civic leaders and popular press, as well as in organizational compensation policies adopted early in the pandemic.\(^2\) Yet, many of the same workers considered ‘essential’ in the fight against COVID-19 are also among the most vulnerable (i.e. part-timers, youth, female, immigrant etc.) in the labour market.

The incongruency between popular support of frontline staff and their labour market realities is what motivated us to look at the relationship between wages and exposure risks to COVID-19. We are interested in whether workers received a compensating differential (CD) based on their exposure risk to contracting the SARS-COV-2 virus. We also investigate what role, if any, unions play in compensating employees most at risk of contracting COVID-19 at work.

We use data from a ranking system developed by Visual Capitalist™ (Lu, 2020) that scored occupations in the US O*NET\(^3\) database for exposure risk to COVID-19 and matched it with Canadian occupational data contained in the Public Use Microdata File (PUMF) of the Labour Force Survey (LFS). From this, we can categorize Canadian occupations based on exposure risk to COVID-19. We find a long-tailed distribution of COVID-19 risk across occupations, with most jobs at the low-to-moderate end of the risk spectrum. Consistent with similar fat-tailed/power law distributions associated with the pandemic (Wong & Collins, 2020), a relatively small number (10%) of occupations account for the majority (90%) of high-risk COVID-19 exposure. Occupations included in this high-risk category are, as expected, jobs in frontline protection, acute care and primary health and retail workers.

With this categorization of COVID-19 exposure risk, we estimate the probability of being in a high-risk occupation based on a set of observable characteristics. We find that workers with more labour market vulnerabilities (e.g. the young, women and immigrants) are more likely to be employed in occupations with high COVID-19 exposure risk. However, due to the ‘location’

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1 Lockdown orders and work from home directives only ever covered about 1/3 of the workforce in most North American jurisdictions, leaving 2/3 exposed to COVID at work (Lu, 2020). This was prefigured in bureau of labour statistics (BLS) force data from 2018. See: https://www.bls.gov/news.release/flex2.101.htm

2 In Canada, there was cross-party support for a wage top-up for all workers deemed essential in the initial lockdown in the spring of 2020. This list included retail, grocery and warehouse workers. But these supports were pulled in the Fall of 2020 once Canada’s universal emergency relief benefit (CERB) scheme ended and a return to employment insurance system was reinstated. Effectively, when there was no longer a need to worry about higher reservation wages keeping frontline workers out of the labour force, the wage top-ups were taken away by employers and the federal government.

3 The Occupational Information Network (O*NET) is developed under the sponsorship of the US Department of Labor/Employment and Training Administration. The primary product is the O*NET database, containing hundreds of standardized and occupation-specific descriptors on almost 1000 occupations covering the entire US economy. Under COVID, it has been used to determine which occupations face the highest risk of exposure to COVID-19. See Lu (2020).
of such occupations (e.g. in sectors like healthcare), we also find positive associations between the likelihood of high-exposure risk and educational attainment, unionization and public sector employment.

Next, we estimate the relationship between COVID-19 exposure risk and hourly earnings utilizing a rich set of conditioning variables that capture a host of individual demographic, human capital and workplace characteristics. While we do not attempt to infer causality from our OLS estimates, we note that individuals working in jobs with high COVID-19 exposure risks were associated with a significant earnings disadvantage (roughly 8% on average) as compared to their observably equivalent lower risk counterparts. When workers in high-risk occupations were covered by a union, however, they enjoyed an earnings premium of roughly 10% as compared to their high-risk, non-unionized counterparts. Furthermore, we find the counterbalancing effects of unionization for high-risk workers to be largest at the lower end of the earnings distribution, for example the bottom 10th percentile of unionized earners in jobs at high risk of COVID exposure receives a roughly 9.5% wage premium compared to the top 90th percentile who do not receive a statistically significant earnings premium. These estimates are in keeping with the compensating wage literature for hazardous work that flourished from the late 1970s up to the early 2000s.

The Canadian context in which this study is undertaken is important to note. The federal government's initial income support response to the COVID-19 economic lockdowns included minor modifications and increases to existing programs, such as Employment Insurance (EI), the Goods and Services Tax credit (GST-C) and the Canada Child Benefit (CCB). Early in the pandemic, the federal government introduced the Canadian Emergency Relief Benefit (CERB), which provided workers who had lost their jobs (or income) with $2000 every 4 weeks, starting on 15 March 2020 for a maximum of 28 weeks (Koebele et al., 2021). Workers who made less than $5000 in 2019 (or during the prior 12 months) were ineligible. The government also (later) introduced the Canada Emergency Student Benefit (CESB), an income support program targeted towards students, as well as a wage supplement for essential workers, negotiated with the provinces. The Canada Emergency Wage Subsidy (CEWS) provided businesses with up to 75% of their employees’ wages (to a maximum amount of $847 per week) — a major increase over the initial announcement of a 10% wage subsidy.4

While the pandemic and subsequent institutional responses are continuing to evolve in Canada and elsewhere, our study has several unique findings and broad implications. First, our study is the first, to our knowledge, to test whether the mechanism of CDs is operative for COVID-related hazardous work, and if so, by how much are workers being compensated for this novel work-related risk in Canada. Second, it is the first article to our knowledge to examine the role of unions in moderating any CD associated with COVID-19 exposure risk. Third, our findings, more generally, update work on CDs largely abandoned two decades ago, which identified conditions under which wages efficiently and equitably compensate workers for risk. Finally, our analysis underscores the important role of institutions such as unions in helping labour markets adjust to emerging dimensions of workplace risk.

4 One important thing to note is that, except for the small top-ups to the GST-C and the CCB, the majority of other individual and household crisis support programs introduced during the pandemic were directly tied to labour market participation (i.e. being employed or having lost a job). Moreover, neither EI nor the CERB allowed workers to collect benefits if they had quit their job for any number of relevant COVID-19–related reasons. A more detailed summary of Canada’s crisis income support programs for workers, individuals and households is summarized in Koebele et al. (2021).
2 | CONCEPTUAL BACKGROUND

2.1 | Compensating wage differentials and workplace risk

Adam Smith⁵ was the first to point out the importance of non-pecuniary aspects of work, such as the disagreeableness or riskiness of a job, in the determination of wages. Since markets can account for such job characteristics by adjusting prices, Smith reasoned that workers would receive wages that, in addition to reflecting their ‘own’ productivity, would also compensate individuals for unpleasant work-related attributes.⁶

Finding empirical evidence of CDs, however, proved challenging. Consider, for example, that any raw estimate of job-related risk has long been associated with lower wages – the exact opposite prediction of compensating wage theory (Robinson, 1988, 1991). This is because CDs operate ceteris paribus (i.e. all other determinants of earnings must be accounted for before we can accurately identify the compensating wage effects of workplace hazards). Moreover, because of the positive income effect associated with the demand for safety (i.e. safe jobs are ‘normal’ goods), the most attractive jobs in society tend to be preferred by workers with higher earning potential; yet, another factor leading to higher pay in less risky jobs (Gunderson & Hyatt, 2001). To disentangle the wage-risk trade-off from other factors that affect wages, economists have relied on models that control for differences in worker productivity as well as different qualitative components of the job and workplace risk (Viscusi & Aldy, 2003; Lavetti, 2020).

Given the conceptual difficulties noted above, perhaps it is not surprising that between Adam Smith’s depiction of the CD mechanism and the first empirical studies testing the theory, there was roughly a 200-year gap. It was not until the mid-1970s that formal ‘models’ emerged detailing the effect of hedonic/qualitative variables on prices and wages (Rosen, 1974; Lucas, 1977; Smith, 1979; Hersch & Viscusi, 1990; Weiss, 1976; Thaler & Rosen, 1975; Sattinger, 1977; Rosen, 1986)⁷ and empirically tested (Brown, 1980; Duncan & Holmlund, 1983; Eberts & Stone, 1985; Garen, 1988; Hersch, 1989; McNabb, 1989; Kostiuk, 1990).⁸

The 20-plus years of published work on estimated trade-offs between wages and workplace risks that began in the late 1970s produced the following empirical regularities (Viscusi & Aldy, 2003):

- The marginal worker, as theory would predict, receives a compensating wage premium when they accept a job with undesirable non-pecuniary characteristics.
- This empirical result emerges in most models that effectively capture workplace risk of injury or mortality and all relevant conditioning variables.⁹

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⁵ As Adam Smith (1776) noted in The Wealth of Nations: ‘The wages of labour vary with the ease or hardship, the cleanliness or dirtiness, the honourableness or dishonourableness of the employment’ (p. 112).

⁶ These non-wage aspects of the job are known as ‘hedonic’ factors in labour economic models. In consumer models, hedonic factors refer to qualitative attributes of products, such as the ‘look’ or ‘social desirability’ of an item or service.

⁷ Some argue that the modelling lineage is cleaner; with Rosen’s (1974) model of hedonic price determination followed by its application by Thaler and Rosen (1975) to risk of injury on the job, being enough to clarify the interaction that ultimately determines compensating payments in competitive markets (Fairris, 1989).

⁸ See Daniel and Sofer (1998:547) for a concise review of this foundational literature.

⁹ There is an interesting exception to this tendency – detailed in the Methods section – whereby the inclusion of union coverage and detailed industry and workplace variables (even standard high-level industry dummies) renders compensating
Extensions of the research findings above have emerged, albeit sporadically, sometimes capturing endogenous risk preferences (i.e. the fact that persons of higher earnings potential will prefer safer work) or likewise capturing self-selection into jobs of varying degrees of risk (i.e. the selection that occurs as workers are sorted into jobs in part based on unobserved factors that affect their productivity, and hence wages). Concerns over self-selection and endogeneity, however, seem to produce contradictory predictions over the ‘accuracy’ of OLS estimation (Lavetti, 2020).

Siebert and Wei (1994) have found that failing to account for the endogeneity of risk can underestimate the risk premium compared to a standard least squares approach. Other research, however, has illustrated the potential for over-estimating the risk premium by failing to control for unobservables. Shogren and Stamland (2002) note that workers with the ability to avoid injury select into risky jobs, while those less able to avoid injury (‘accident prone’ workers) select into less-risky jobs (implying that risk premiums could be grossly overestimated). More recent estimates by Guardado and Zeibarth (2019) confirm the importance of controlling for unobserved time-invariant worker heterogeneity, either through a richer set of covariates or panel data.

With that background in mind, this article focuses on a subset of the CD for hazardous work literature that began in the early 1980s; namely, the study of differences between union and non-union compensation for exposure to job/workplace risk.

2.2 The ‘moderating’ role of unions in compensating for hazardous work

The CD literature has consistently found that union members receive much larger hazard pay premiums than non-union workers (Dickens, 1984; Fairris, 1989 & 1992; Seibert & Wei, 1994; Moore, 1995). In fact, the strongest finding is the presence, even in richly specified models, of negative CDs (i.e. relatively high risk and low wages) for non-union workers and large positive CDs for union workers (Seibert & Wei, 1994; Dorman & Hagstrom, 1998). It turns out that the mechanisms generating compensating wage differentials and their magnitude are qualitatively different in the union and non-union sectors of the economy (Fairris, 1989, 1992).

Dickens (1984) was the first to summarize the literature (five articles at that time) on union versus non-union differences in CDs. He found that Thaler and Rosen (1975) were the first to report on the union versus non-union difference. In all four of their wage equations, they estimated that a union interaction with risk of fatal injury was significantly positive, while the risk term was not. Estimated CDs were 8 to 10 times greater for union than non-union workers. This higher CD for risk among union workers was evident in all articles reviewed at the time (see Viscusi, 1980; Olson, 1981; Freeman & Medoff, 1981; and Dorsey, 1983).

Articles published subsequently (see Fairris, 1989; Fairris, 1992; Seibert & Wei, 1994; Moore, 1995; Daniel & Sofer, 1998; Dorman & Hagstrom, 1998), often incorporating larger datasets, endogenous risk preferences and self-selection modelling, have only served to reinforce the findings from the foundational literature. In fact, not only are union settings associated with larger compensating payments for hazardous work, but several studies corroborate negative CDs – relatively high risk and low wages – for non-union workers (Fairris, 1992; Dorman & Hagstrom, 1998). Moreover, the earlier work which estimated differences between risk premiums in union and non-union wages without accounting for possible self-selection may have underestimated the positive wages for workplace hazards insignificant and, in some cases, negative (see Leigh, 1991; 1995 and Dorman and Hagstrom 1998).
union effect. When job risk is made endogenous to avoid a selectivity bias arising if more able people choose safer jobs, Seibert and Wei (1994) found that this adjustment has a considerable positive effect on the union group, while lowering that of non-union workers.

The above results paint a remarkably consistent portrait that CDs associated with workplace hazards are larger for union workers and that non-union workers possibly receive negative CDs. The possible mechanisms are varied and have been detailed in greater depth elsewhere (Dickens, 1984; Fairris, 1992; Moore, 1995; Dorman & Hagstrom, 1998). Below, we highlight a few of the most obvious channels.

Some argue that positive CDs for hazardous work in union settings are due to the insensitivity of competitive firms to the preferences of inframarginal workers (Viscusi, 1980). One interpretation of such a channel is that collective bargaining acts as an ‘efficient’ institution, in the sense that it reveals worker preferences to management, which in firms lacking unions, leads to an under-provision of safety in competitive (non-union) firms. Olson (1981) has noted that larger union CDs likely reflect unions’ informational advantages (i.e. unions are better at detecting latent workplace risk and at informing members and management). Likewise, a negative CD for non-union workers is viewed by Moore (1995) as evidence of worker-specific, or supply-side risk that is removed by the presence of a union, based on the belief that unionized firms will be more likely to filter out high-risk unproductive workers. Once again, the implication would be that occupational health and safety (OHS) is underprovided in the non-union sector.

There are also simple bargaining models that emphasize the ‘institutional’ differences between the sectors. In the non-union sector, competitive labor markets establish a set of ‘implicit prices’ for bad working conditions that are more reflective of the preferences of individual workers and the technologies of individual firms. In the union sector, compensating payments are the outcomes of collective bargaining agreements, in which collective preferences and collective worker power are important factors (Dickens, 1984). Union and non-union sectors will, therefore, generate different risk-wage trade-offs. Indeed, a study of German workers found that there is evidence of a positive relationship between risk aversion and the likelihood of union membership for full-time employees (Goerke & Pannenberg, 2012).

Finally, larger compensating wages may reflect a union ‘voice’ effect, in which unions negotiate compensation for hazardous work that otherwise does not emerge via ‘exit’ forces (i.e. greater job vacancy or employee quits) alone (Fairris, 1989, 1992). Unions can achieve these gains by protecting workers from reprisal if they exercise their right-to-refuse unsafe work, a protection rarely enforced in non-union settings (Sojourner & Yang, 2015).11

Regardless of the channel, revisiting the literature on CDs and applying it to the emergence of new workplace risk (in the form of COVID-19) is expected to generate CDs that differ significantly between union and non-union workers, with union members receiving a larger CD than (observably equivalent) non-union workers. The policy implications of such a finding, if found, will be discussed in our conclusion.

Although our focus is on compensating wage differentials, it bears mentioning that unions play an important role in all aspects of OHS. Unions have, for example, shaped the development and ensured the enforcement of OHS laws (e.g. Mischel & Walters, 2003), influenced workers’ perceptions of safety (Sinclair et al., 2010), communicated important safety information (e.g. Gillen

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10 Interestingly, allowing for simultaneity in the relationship between earnings and injuries has only a slight effect on the OLS results in most papers. There is also no significant evidence of simultaneity bias in the earnings coefficients of OLS estimated injury rate equations in either the union or non-union sector (Fairris, 1992).

11 We would like to thank an anonymous reviewer for this point.
et al., 2002), facilitated the transfer of OHS knowledge through training (Hilyer et al., 2000) and empowered workers to exercise rights under OHS legislation (Weil, 1992). It is expected, therefore, that unions would be key stakeholders in the present health crisis, which includes, but is certainly not limited to, negotiating for higher compensating wage premiums particularly for those workers most at-risk of exposure to COVID-19.

2.3 Main research questions

Based on the review of literature and the nature of the current COVID crisis, we seek to answer the following six questions: (1) What characteristics are significantly associated with the probability of employment in an occupation with a high COVID-19 exposure risk? (2) Did workers receive a pandemic pay premium during the first COVID-19 lockdown and was the pandemic pay premium larger (smaller) for those at greater risk of contacting COVID-19 (i.e. is there evidence of the ‘frontline hero pay’ effect?) (3) Do workers in high COVID-19 risk occupations earn more (less) than observably similar workers in lower COVID-19 risk occupations? (4) Does union coverage (positively) moderate the pandemic pay premium for workers in occupations facing high COVID-19 exposure risk, as past research would predict? (5) Does this (mean) estimate of union CDs for COVID-19 exposure risk differ by sub-groups (male vs. female) and across the earnings distribution? And (6) How does the addition of a standard lost-time injury measure of workplace risk alter the CD estimates of Covid-related risk?

3 DATA

The data for the present analysis are obtained from a combination of sources. We first employ the Public Use Microdata Files (PUMF) of the Canadian LFS for the months of March–September inclusive in 2020 and 2019. The LFS is conducted by Statistics Canada on a monthly basis and includes responses from private Canadian households in all 10 provinces across the country. The LFS is a cross-sectional survey that samples using a 6-month rotating panel (Statistics Canada, 2016). Since we are interested in observing wages during the COVID-19 pandemic, we use all seven consecutive months (March–September) in 2020 and the same months in 2019 as the pre-COVID-19 reference period. Our sample includes employed persons working part-time or full-time between the ages of 20 and 64 years, reporting positive (non-missing) hourly earnings. Self-employed respondents, those currently in school and multiple job holders are excluded from the analysis. The hourly earnings variable in the LFS is constructed in combination with the usual hours of paid work per week (Statistics Canada, 2016, 19). In this question, ‘respondents are asked to report their wage/salary before taxes and other deductions, and include tips and commissions during the reference week (Statistics Canada, 2016, 19). Furthermore, ‘the survey or collection period, which is ten days following the reference week (Sunday to Tuesday), is when the LFS interviews are conducted’ (Statistics Canada, 2016, 16).

12 By using this period, this means that some of the same respondents will be sampled more than once and up to six times. To determine the possible influence of repeated measures, we re-estimated the analysis using only March and September. The results are qualitatively similar and are available from the authors upon request.

13 We excluded youth from our analysis for an obvious reason. During the height of the pandemic in early Spring 2020, the unemployment rate for youth (aged 15–19) soared to close to 30%. Such a high drop off in employed youth was noticeable in the 2020 sample, creating the potential for sample selection bias between the 2019 and 2020 periods.
These sample restrictions give us a total of 536,324 observations.\textsuperscript{14} This is our ‘pooled’ sample size that includes two distinct cohorts of workers in the 2019 and 2020 LFS survey waves, which are distributed almost evenly in our sample (51% and 49%, respectively).\textsuperscript{15}

3.1 Measuring COVID-19 exposure risk by occupational characteristics

Our first objective is to classify occupations according to their risk of contracting COVID-19. We take an approach inspired by Tiagi (2015) and we triangulate data from the aforementioned LFS data with a system developed by Visual Capitalist\textsuperscript{SM} (Lu, 2020) that scores occupations based on COVID-19 exposure risk. The score uses characteristics found on the US Department of Labor’s Occupational Information Network (O*NET) database and is based on three dimensions (Lu, 2020):

1. **Contact with others at work:** How much does this job require the worker to be in contact with others in order to perform it on a daily basis?
2. **Physical proximity to others at work:** ‘To what extent does this job require the worker to perform tasks in close physical proximity to others on a daily basis?’
3. **Exposure to disease and infection at work:** How often does this job require exposure to hazardous conditions on a daily basis?’

Scoring each component above out of 100 and weighting equally generates an overall ‘COVID-19 Risk Score’ between 0 and 100 is generated, where 100 represents the highest possible risk of COVID-19 exposure (Lu, 2020). By applying this methodology to the occupational characteristics data, Visual Capitalist\textsuperscript{SM} provided COVID risk scores for 966 occupations in the O*Net database (Lu, 2020). Using the Canadian National Occupational Classification (NOC) system, we match the US occupations onto Canadian ones and assign the COVID-19 exposure risk scores that correspond to their US counterparts. The details of this matching and variable construction process can be found in online Appendix A1; however, the distribution of COVID-19 risk scores for the 40 occupational categories available in the LFS is plotted in Figure 1.

We assign occupations in the 90th percentile or higher of the risk score distribution (i.e. a COVID-19 risk score of 40 or higher) to the category of ‘high-risk’ (Level 5) occupations, the remaining occupations are grouped as low-risk (Levels 1–4) occupations. The 90th percentile is not an arbitrary cut-off as it falls where there is a natural break at 40 in the distribution of scores. The occupational risk scores then jump up again (dotted red line in Figure 1) with almost all ‘high-risk’ occupations scoring between 60 and 80. Online Table A2 details the major occupational groups in each risk level and corresponding COVID-19 risk scores.

The COVID risk exposure index is a prospective measure, based on factors that were known to be associated with coronavirus transmission in the first phase of the pandemic. In studies conducted retrospectively, occupations identified as having the highest potential risk also turned out to be the occupations with the greatest risk of contracting the virus (Nguyen et al., 2020).

\textsuperscript{14} Bearing in mind footnote 13, given the sampling structure of the LFS, some respondents will be included up to six times in the data set.

\textsuperscript{15} The 2-percentage point drop in the employed LFS sample between 2019 and 2020 reflects the drop off in employed persons in the LFS survey due to COVID lockdown measures, which shuttered large parts of the service and retail economy.
METHODS AND MODEL SPECIFICATIONS

Given the classification of occupations based on COVID-19 exposure risk, we next describe our estimating strategy for measuring the difference in earnings across individual workers exposed to differing levels of COVID risk, as well as the extent to which unions influence these compensating wage differentials before and during the COVID pandemic.

4.1 Estimating the probability of being at high risk of contracting COVID

We begin with a probability model estimating the likelihood of being employed in an occupation with a high risk of contracting COVID-19 (individual subscripts supressed):

$$\Pr (\text{Highcovrisk}) = \alpha + \beta_1 X + \beta_2 Union + \varepsilon$$  (1)

Equation (1) is estimated for the entire sample of employed persons and then separately for unionized and non-unionized workers.\(^\text{16}\) The outcome $\text{Highcovrisk}$ is a dichotomous variable denoting work in an occupation with a high COVID-19-exposure risk score, coded ‘1’ if the

\(^{16}\) Models are estimated with a linear probability (OLS) regression model. The coefficients which show the percentage point change in probability at the mean value of the regressors (average marginal effects) and t-statistics are obtained from STATA’s post-estimation ‘margins’ command. We repeat the analysis using logistic regression and obtain nearly identical
respondent works in such a job with a risk score > 40 and ‘0’ otherwise.\textsuperscript{17} Except for several interaction terms omitted from the probability estimations, the vector of covariates denoted by \( X \) are the same as those used in the OLS earnings regressions (described in Section 4.2). Our expectation, based on past research, is that the coefficient \( \beta_2 \) for our \textit{Union} variable – which is coded ‘1’ if a respondent is covered by a collective agreement and ‘0’ otherwise – will be positive, indicating that union members are more likely employed in high-risk jobs (Fairris, 1992).

There are a variety of reasons for this positive union-risk correlation, including the obvious channel whereby worker demand for union representation rises with workplace risk. There is also the complicated ‘bumping’ and ‘bidding’ rights in union contracts that may make unionized workers less safe in production, since labor allocation often depends on seniority rules rather than ability (Appleton & Baker, 1984). Alternatively, unionized employers, responding to the supra-competitive wages won in union contracts, may worsen workplace safety in an attempt to offset union rents and maintain profits (Duncan & Stafford, 1980).

4.2 Estimating the compensation wage differential associated with COVID-19 risk

Next, we present our estimating strategy for the earnings equations, where the natural logarithm of hourly earnings \( W \) (in constant March 2019 dollars) is regressed on a number of observable characteristics (individual subscripts suppressed), with a standard error \( \varepsilon \) term capturing random variation:

\[
\ln W = \alpha + \beta_1 X + \beta_2 \text{Union} + \beta_3 \text{Highcovrisk} + \beta_4 \text{Covidperiod} \\
+ \beta_5 (\text{Highcovrisk} \times \text{Covidperiod}) + \beta_6 (\text{Highcovrisk} \times \text{Union}) \\
+ \beta_7 (\text{Union} \times \text{Covidperiod}) + \beta_8 (\text{Highcovrisk} \times \text{Union} \times \text{Covidperiod}) + \varepsilon,
\]

where \( X \) is a vector of determinants of individual earnings most often used in prominent risk-wage studies, that is: (i) \textit{core demographic and human capital variables}, such as sex, level of educational attainment, age, marital status, immigrant status, the presence of own children under 12 years of age, geographic region and living in a large urban area; and (ii) \textit{employer and job characteristics}, such as working part-time, tenure in the current job (years), employment in

\textsuperscript{17} We also ran the models with alternate specifications of COVID-19 exposure risk: one where COVID-19 risk is a continuous variable ranging in values from 15.06 to 76.38 and a second specification that uses five groupings of COVID-19 occupational risk, with group 5 being the highest risk group and equivalent to the high COVID-19 risk variable used here. Notably, in the specification with five risk groups, risk group 4 experienced the largest wage penalties, followed by risk group 5; however, risk group 5 enjoyed the largest union wage premiums. The results of these models are available from the authors upon request.
the public sector and firm size (measured in number of employees). Our preferred model specification adds seasonal dummies (capturing seasonality effects) and indicators for (i) and (ii) to Equation (2). The most saturated model adds several broad industry categories. The inclusion of industry dummies in CD equations, as noted in Section 4.3, has been a source of contention in past modelling.18

We also include in Equation (2) our key explanatory variables denoting whether a respondent is covered by a collective agreement Union and whether they work in an occupation with a high risk of contracting COVID-19 (Highcovrisk). The variable Covid period is a dummy capturing the months of March–September 2020 when various COVID lockdown orders and restrictions were introduced, the omitted reference being the same months in 2019 prior to the pandemic. To capture the differential effect of the COVID-19 crisis on different groups, we include an interaction term that covers being in a high-risk occupation during the initial COVID-19 period (Highcovrisk *Covidperiod); an interaction term estimating the differential earnings impact of unionization for workers in occupations with high-COVID risk (Highcovrisk *Union); a similar term capturing the differential earnings impact of unionization for workers employed during the first 6 months of COVID period (Union * Covidperiod); and finally, a three-way interaction term combining unionization, the COVID-19 time period and working in an occupation with a high risk of COVID exposure (Highcovrisk *Union * Covidperiod).

Lastly, using the same specification noted in model 3, which includes the interaction terms, we examine the relationship between risk of exposure to COVID-19 and earnings across the earnings distribution. To do this, we employ recentred influence functions (RIFs), a type of unconditional quantile regression developed by Firpo et al. (2007, 2009). The RIF technique is an unconditional estimate which means that unlike traditional quantile regression, which is conditional and holds constant the mean value of the other covariates in each quantile, the unconditional RIF holds constant the entire distribution of covariates. The use of a distributional approach such as RIF, therefore, allows us to observe the impact of a one-unit change in a variable of interest, in our case – being in a high COVID-19 risk occupation – and earnings at different points on the earnings distribution (e.g. Firpo et al., 2009).

What does this mean in a practical sense? It is quite conceivable that low (high) wage workers experience different compensating premia (penalties) associated with high COVID-19 risk work and that, further, unions may have a differential impact not only for all workers at different points along the earnings function, as has been demonstrated in earlier research (e.g. Firpo et al., 2009), but also distinctly advantage (disadvantage) high COVID-19 risk workers at various points on the earnings distribution.

18 The issue is particularly front and centre in Dorman and Hagstrom’s (1998) critique of the competitive labour market model as the basis for estimating returns to risk. The authors highlight a common finding in US studies, that is when industry (and occupation) dummy variables are included in the wage equation, the coefficients on the risk variables are smaller and/or their statistical significance is diminished relative to otherwise identical regressions without industry dummies. Dorman and Hagstrom (1998) argue that this calls into question the exclusion of industry dummy variables in the wage equations. In our case, given the controversy noted above and given the potentially high level of collinearity between occupational groupings, industry and employment in the public sector, we follow Gunderson and Hyatt (2001), in that we have used industry dummies, but their inclusion is merely to benchmark other studies and do not reflect our preferred specification.
4.3 Compensating differentials for risk versus inter-industry wage differentials

An interesting sideline to the empirical literature on compensating wage differentials (CDs) for hazardous work is the role played by industry and workplace-level variables (Viscusi & Aldy, 2003) since their inclusion in estimating wage equations may reflect industry/workplace wage premiums because fatality/injury risk variables typically reflect industry-level rather than individual risk (Leigh, 1995; Dorman & Hagstrom, 1998). Both Leigh (1995) and Dorman and Hagstrom (1998) contend that the early CD literature conveniently left out estimates with industry dummies. Their own estimates show insignificant CD coefficients for workplace risk when a rich set of industry and workplace variables are added; a finding they take to mean the failure of market mechanisms to adequately compensate workers for risk.

Viscusi and Aldy (2003) dispute these conclusions and contend that the earlier literature cited by Dorman and Hagstrom was incomplete and contains several studies where positive CDs are present with industry dummies (e.g. Viscusi, 1978; Freeman & Medoff, 1981; Marin & Psacharopoulos, 1982; Cousineau et al., 1992). They also contend that inserting industry dummy variables into the CD regression equation induces multicollinearity with the risk variable (Hamermesh and Wolfe (1990)).

What the debate over industry-level variables appears to illustrate is that risk measures associated with an industry, occupation or job, rather than an individual, can act as collinear regressors in estimates with too few observations. Adding a rich set of industry variables in a model with an industry-derived workplace hazard will, in small samples, necessarily increase standard errors and possibly dampen the magnitudes of the parameter estimates. We, therefore, present estimates with industry dummies in column 4 of our Table 3 but base our discussion on models that exclude detailed industry dummies in our wage regression tables, repeating the analysis separately for males and females and across the earnings distribution using the non-industry specification.

4.4 Sample means for key variables and controls

Table 1 provides means for the entire sample and then separately for high and low COVID-19-risk workers. Noteworthy is the proportion of the sample who work in high COVID-risk occupations, which is 13% across the entire sample, but markedly higher (21%) for workers covered by a collective agreement. Women and younger workers are also more highly represented in high-risk occupations. Interestingly, up to the point of a bachelor’s degree, high-risk occupations appear to have a lower proportion of workers with high school or less as their highest level of education. However, the proportion of workers with the highest level of education, above a bachelor’s degree, is greater in occupations with low COVID-exposure risk. COVID-19 exposure risk is higher in the public sector and for those working in part-time jobs, where there is a doubling of the probability of being in a high-risk occupation (18% vs. 9%). Finally, the raw sample means show lower wages for workers in occupations with high COVID exposure risks as compared to those in lower risk occupations ($28.92 vs. $30.59) – evidence of a negative CD for COVID exposure risk – and higher wages for those covered by a union collective agreement ($33.03 vs. $29.11).
### TABLE 1 Sample means for estimating variables

|                          | Whole sample [1] | By COVID risk | By union coverage |
|--------------------------|------------------|---------------|------------------|
|                          | High [2]        | Low [3]       | Yes [4]          | No [5]          |
| Real wage ($)            | 30.38           | 28.92         | 30.59            | 33.03           | 29.11           |
| Ln (Real wage)           | 3.41            | 3.36          | 3.42             | 3.49            | 3.37            |
| Covid risk score (0–100) | 30.85           | 59.91         | 26.66            | 34.96           | 28.89           |
| [Covid risk low-to-moderate] | 0.87      | –             | –                | 0.79            | 0.91            |
| Covid risk high          | 0.13            | –             | –                | 0.21            | 0.09            |
| [Non-union]              | 0.68            | 0.47          | 0.71             | –              | –               |
| Union                    | 0.32            | 0.53          | 0.29             | –              | –               |
| [Pre-Covid period Mar–Sep 2019] | 0.51     | 0.52          | 0.51             | 0.50            | 0.52            |
| Covid period Mar–Sep 2020 | 0.49         | 0.48          | 0.49             | 0.50            | 0.48            |
| **Demographic/human capital** |               |               |                  |                 |
| [Male]                   | 0.52            | 0.26          | 0.56             | 0.48            | 0.54            |
| Female                   | 0.48            | 0.74          | 0.44             | 0.52            | 0.46            |
| [< High school]          | 0.06            | 0.03          | 0.06             | 0.04            | 0.06            |
| High school              | 0.17            | 0.12          | 0.18             | 0.13            | 0.19            |
| Post-secondary certificate | 0.43         | 0.51          | 0.42             | 0.46            | 0.41            |
| Bachelors                | 0.24            | 0.26          | 0.24             | 0.25            | 0.23            |
| Post-graduate            | 0.10            | 0.08          | 0.11             | 0.11            | 0.10            |
| [Age 20–29]              | 0.21            | 0.25          | 0.20             | 0.15            | 0.24            |
| Age 30–59                | 0.72            | 0.69          | 0.72             | 0.77            | 0.69            |
| Age 60–64                | 0.07            | 0.06          | 0.07             | 0.07            | 0.07            |
| [Single]                 | 0.27            | 0.28          | 0.27             | 0.22            | 0.29            |
| Divorced/widowed         | 0.07            | 0.08          | 0.07             | 0.08            | 0.07            |
| Married                  | 0.66            | 0.65          | 0.66             | 0.70            | 0.64            |
| [Non-immigrant]          | 0.75            | 0.74          | 0.75             | 0.80            | 0.73            |
| Immigrant                | 0.25            | 0.26          | 0.25             | 0.20            | 0.27            |
| [No children]            | 0.72            | 0.71          | 0.73             | 0.70            | 0.74            |
| Children in household    | 0.28            | 0.29          | 0.27             | 0.30            | 0.26            |
| [Rural/small urban]      | 0.44            | 0.47          | 0.43             | 0.48            | 0.41            |
| Large urban centre       | 0.56            | 0.53          | 0.57             | 0.52            | 0.59            |
| **Job/workplace controls** |               |               |                  |                 |
| [Full-time]              | 0.90            | 0.82          | 0.91             | 0.91            | 0.90            |
| Part-time                | 0.10            | 0.18          | 0.09             | 0.09            | 0.10            |
| Tenure (years)           | 7.53            | 7.71          | 7.51             | 9.69            | 6.51            |
| [Private sector]         | 0.75            | 0.54          | 0.78             | 0.39            | 0.92            |
| Public sector            | 0.25            | 0.46          | 0.22             | 0.61            | 0.08            |
| [Firm size < 20 employees] | 0.17       | 0.13          | 0.17             | 0.04            | 0.23            |
| Firm size 20–500 employees | 0.31       | 0.27          | 0.33             | 0.23            | 0.36            |
| Firm size 500+           | 0.52            | 0.59          | 0.50             | 0.74            | 0.41            |
| Observations             | 536,324         | 70,688        | 465,636          | 190,838         | 345,486         |

Notes: Excluded reference categories are denoted by [ ] and italics. All variables are categorical (0,1) measures unless denoted by measured values in (). Additional covariates not shown here include region/province of residence and survey month. Dummy variable sets may not sum to 1 due to rounding.
5 | RESULTS

5.1 | The probability of being employed in an occupation with high COVID-19 exposure

Table 2, column 1, uses the whole sample to estimate the probability of working in an occupation with a high risk of COVID-19 exposure. That probability is 0.126 (or roughly 12.6%). Unionized workers are 0.074 (about 7.4 percentage points) more likely than non-unionized workers to be in a high COVID-19 exposure risk occupation (Table 2, column 1, row 2). In terms of the relative risk, which is based on the baseline probability of 0.126, the union exposure risk is roughly 60% higher than for non-union workers. Looking at characteristics typically associated with historical market disadvantage, we see that females (roughly 10.6 percentage points or 84%), immigrants (roughly 3.7 percentage points or 30%) and younger workers (5.4 and 7.3 percentage points or 40% and 58% more than 30- to 59- and 60- to 64-year-olds, respectively) are all significantly more likely to be employed in occupations at high risk of contracting COVID-19. We also find a strong positive association between part-time employment and the probability of working in an occupation with a high COVID-19 exposure risk – part-time workers are about 8.7 percentage points (or 70%) more likely to be in a high-COVID risk job than a full-time worker.

Interestingly, compared to those with less than high school, all levels of education at or below a bachelor’s degree are positively associated with a greater probability of employment in a high-risk occupation and it is not until one has a post-graduate or graduate degree (i.e. education above a bachelor’s degree) where we see a negative and statistically significant association. This inverted-U result likely reflects the skilled occupations in health care and frontline public protection that make up a large portion of the jobs in high-COVID-risk categories. The negative relationship between post-graduate credentials and high-risk work is also indicative of a ‘white-collar’ divide at the top echelon of occupational COVID-19 risk.

Other relationships, which are non-significant in the whole sample estimates (Table 2, column 1), show significantly divergent patterns between the union (Table 2, column 2) and non-union samples (Table 2, column 3). For example, the dummy variable capturing work during the first 6 months of the COVID pandemic (March–September 2020) shows a small but significant increase of about 0.9 percentage points for those covered by a union (which relative to the overall probability of 12.6% represents a 6% greater risk of COVID exposure than the same period in 2019). By way of contrast, there is an almost equivalent decline of –0.6 percentage points (or 5% lower probability) for non-union workers.

This pattern is repeated when we look at firm size. For unionized workers, larger firm size is associated with greater COVID exposure risk – likely reflecting large employers in the health sector, postal offices and UPS workers, long-term care homes organized by unions, such as SEIU, and a non-trivial number of large food retailers organized by UFCW in Canada. For non-union workers, however, the pattern is reversed, reflective of large non-union employers in finance, insurance and IT that were at low risk prior to the pandemic (based on the O*NET criteria) and became safer still once white-collar employees could work from home.

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19 We recognize that coefficients for dummy variables in OLS models are not exactly equivalent to percentages (Halvorsen & Palmquist, 1980). In log earnings equations, the true proportional change is \( \exp(\hat{\beta}) - 1 \), where \( \hat{\beta} \) is the estimated coefficient. For low values of \( \hat{\beta} \), the approximation is very close, underestimating the true value by less than 0.02 for values of \( \hat{\beta} < 0.20 \), which is the case for our main variables of interest. Therefore, for ease of interpretation use, refer to the OLS regression coefficients as approximate percentages.
### Table 2: Probability of being employed in an occupation with high risk of COVID exposure

| Dependent variable | Whole sample dy/dx (t-stat) | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|--------------------|-----------------------------|----------------------|--------------------------|
| Mean Prob. High Risk | 0.126** (211.01) | 0.206** (167.50) | 0.088** (134.93) |

**Independent variables**

| [Non-union] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-------------|----------------------|--------------------------|
| Union | 0.075** (44.57) | – | – |

| [Pre-Covid period] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|--------------------|----------------------|--------------------------|
| Covid period | –0.001 (−0.71) | 0.009** (3.79) | −0.006** (−4.45) |

| [Male] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|--------|----------------------|--------------------------|
| Female | 0.106** (83.37) | 0.147** (53.75) | 0.087** (62.66) |

| [Non-immigrant] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------------|----------------------|--------------------------|
| Immigrant | 0.036** (21.83) | 0.082** (22.25) | 0.016** (9.28) |

| [Age 20–29] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|--------------|----------------------|--------------------------|
| Age 30–59 | −0.045** (−23.61) | −0.096** (−21.87) | −0.025** (−12.41) |
| Age 60–64 | −0.062** (−22.25) | −0.123** (−19.94) | −0.037** (−12.58) |

| [Less than high school] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-------------------------|----------------------|--------------------------|
| High school graduate | 0.010** (4.46) | 0.000 (0.07) | 0.015** (5.76) |
| Some post-secondary | 0.018** (5.32) | 0.021** (2.95) | 0.018** (4.93) |
| Post-secondary certificate | 0.059** (27.92) | 0.120** (26.57) | 0.033** (13.90) |
| Bachelor’s degree (BA) | 0.022** (9.41) | 0.058** (11.09) | 0.007** (2.71) |
| Above BA | −0.017** (−6.15) | −0.041** (−7.25) | 0.004 (1.25) |

| [Single] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------|----------------------|--------------------------|
| Married | −0.010** (−5.82) | −0.004 (−1.06) | −0.013** (−7.20) |
| Divorced/widowed | −0.007** (−2.60) | 0.003 (0.48) | −0.013** (−4.52) |

| [No child 0–12 years] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------------------|----------------------|--------------------------|
| Pres. of child 0–12 years | 0.009** (6.35) | 0.021** (7.07) | 0.003 (1.80) |

| [Rural or small urban] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------------------|----------------------|--------------------------|
| Large urban area | −0.008** (−6.09) | −0.008** (−3.22) | −0.007** (−5.16) |

| [Full-time work] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|------------------|----------------------|--------------------------|
| Part-time work | 0.085** (32.88) | 0.108** (21.07) | 0.072** (24.52) |
| Tenure (years) × 100 | 0.018 (1.96) | 0.113** (5.94) | −0.067** (−6.59) |

| [Private sector] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------------|----------------------|--------------------------|
| Public sector | 0.079** (39.33) | 0.085** (28.07) | 0.049** (16.53) |

| [Firm < 20 employees] | Union dy/dx (t-stat) | Non-union dy/dx (t-stat) |
|-----------------------|----------------------|--------------------------|
| Firm 20–500 employee | 0.003 (1.86) | 0.054** (9.65) | −0.003 (−1.87) |
| Firm > 500 employees | 0.001 (0.48) | 0.045** (8.13) | −0.004* (−2.15) |
| Constant | 0.027** (8.40) | 0.002 (0.30) | 0.057** (16.75) |

| Observations | 536,324 | 190,838 | 345,486 |

Significance denoted by * $p < 0.05$, ** $p < 0.01$. Notes: The outcome variable is coded ‘1’ if the respondent works in an occupation with high risk of COVID exposure and ‘0’ otherwise. Models estimated with linear probability (OLS) regression. The coefficients and $t$-statistics obtained from STATA’s post-estimation ‘margins’ command. Models are weighted using STATA’s probability weights (p-weights) and the LFS sampling weight (FINALWT) rescaled for 14 months of data. Sample sizes represent the unweighted count of respondents in the dataset. Additional covariates not shown here include province/region of residence and survey month to capture seasonal variation. Due to small magnitude sizes, the coefficient on the tenure_years (tenure in years) variable is multiplied by 100, unlike in the earnings equations where tenure is expressed in months.
In sum, except for education, we find that traditionally more vulnerable workers – such as females, youth, immigrants and part-time employees – are more likely to be employed in occupations with a high risk of being exposed to COVID-19. This is evident in Figure 2, where the adjusted risks of COVID-exposure risks are shown for each of these vulnerable groups. These findings are, therefore, consistent with the public rhetoric and medical evidence thus far, indicating that the pandemic exacerbates existing economic inequalities, often disproportionately affecting marginalized groups in the workplace (Kirby, 2020). Union coverage was also associated with a higher COVID-19 exposure risk, consistent with earlier work showing the positive association between unions and more hazardous work; but also, because in the case of COVID, many non-union employees in the private sector (in particular in sectors like finance and IT) were at lower risk of contracting any virus at work prior to March 2020, and became even less at risk when they were told to work from home.

5.2 Hourly earnings estimates and pay premiums for COVID-19 exposure risk

Table 3 presents OLS estimates from the regression of the logarithm of the hourly wage on our key set of independent variables, the Highcovidrisk variable, and its interaction with the Covidperiod and Union membership dummy variables. The control variables (not reported in Table 3 but the full model output is available from the authors upon request) perform as expected. As reported in our preferred specification in Table 3 (column 3), and consistent with the results in Table 2, for our vulnerable worker populations, significantly lower wages are observed among female, young (aged 20–29), immigrant and part-time workers. These results are remarkably consistent across all specifications, including model 4, which includes industry dummies.
TABLE 3  Estimated log hourly earnings equations and COVID-19 exposure risk

| DV: Ln hourly wages | Ordinary least squares regression |
|---------------------|-----------------------------------|
|                     | [1] | [2] | [3] | [4] |
| **Key independent variables** | | | | |
| [Non-union] | | | | |
| Union | 0.171** | 0.133** | 0.004 | −0.002 |
| (69.91) | (62.10) | (1.53) | (0.79) |
| [Pre-Covid period] | | | | |
| Covid period | 0.075** | 0.058** | 0.053** | 0.051** |
| (31.66) | (28.96) | (27.92) | (27.38) |
| [Covid risk low] | | | | |
| Covid risk high | −0.172** | −0.096** | −0.076** | −0.051** |
| (38.37) | (22.60) | (19.11) | (12.50) |
| Covid risk high × Union | 0.168** | 0.136** | 0.103** | 0.085** |
| (27.91) | (24.67) | (20.18) | (16.46) |
| Covid risk high × Covid period | −0.009 | −0.002 | −0.005 | −0.005 |
| (1.3) | (0.36) | (0.75) | (0.72) |
| Covid period × Union | −0.031** | −0.027** | −0.027** | −0.023** |
| (8.45) | (8.50) | (8.78) | (7.82) |
| Covid risk high × Covid period × Union | 0.034** | 0.026** | 0.020** | 0.020** |
| (3.70) | (3.09) | (2.59) | (2.58) |
| [Male] | | | | |
| Female | −0.163** | −0.158** | −0.135** | |
| (106.58) | (107.17) | (88.38) | |
| [Age 20–29] | | | | |
| Age 30–59 | 0.205** | 0.125** | 0.117** | |
| (101.83) | (60.70) | (58.68) | |
| Age 60–64 | 0.148** | 0.057** | 0.051** | |
| (43.9) | (17.03) | (15.76) | |
| [Non-immigrant] | | | | |
| Immigrant | −0.175** | −0.143** | −0.14** | |
| (84.82) | (72.96) | (73.03) | |
| [Full-time job] | | | | |
| Part-time job | −0.18** | −0.139** | | |
| (74.87) | (59.12) | | |
| Demographic/human capital controls | No | Yes | Yes | Yes |
| Job/workplace controls | No | No | Yes | Yes |
| Industry controls | No | No | No | Yes |
| Constant | 3.232** | 2.82** | 2.766** | 2.601** |
| (2102.29) | (743.33) | (703.79) | (597.53) |
| R-squared | 0.05 | 0.30 | 0.37 | 0.4 |
| Observations | 536,324 | 536,324 | 536,324 | 536,324 |

Significance denoted by * p < 0.05 ** p < 0.01.
Notes: All models estimated using OLS. The outcome variable is the natural logarithm of hourly earnings adjusted for inflation and t-statistics in parentheses. Models are weighted using STATA’s probability weights (p-weights) and the LFS sampling weight (FINALWT) rescaled for 14 months of data. Sample sizes represent the unweighted count of respondents in the dataset. Additional covariates not shown here include marital status, presence of children, region of residence, living in a large urban centre, survey month, tenure at current job (years) and firm size.
5.3 Union coverage and estimated hourly earnings

Being covered by a union is associated with a ‘raw’ 17.1% wage premium in our estimates (Table 3, model 1) when no control variables other than the COVID risk measures and their interactions are added. When individual demographic and human capital controls are added to the wage equation, the union wage premium falls to about 13% (Table 3, column 2) and when the preferred set of workplace and job controls are added (Table 3, column 3), the union wage premium is no longer statistically significant.

This effective ‘0’ union wage premium in a model richly specified with right-hand side conditioning variables is in line with models estimated using panel data and/or selection effects. Moreover, recent estimates in Canada (Campolieti, 2018; Gomez & Lamb, 2019) and elsewhere (Blanchflower & Bryson, 2010) have documented the decline of the union wage premium for the average worker. However, that same research has shown is that the union wage premium differs considerably among occupations (Blanchflower & Bryson, 2010) and by sub-groups, especially among workers traditionally disadvantaged in the workplace (Gomez & Lamb, 2016). It appears that workers who already have substantial labour market leverage benefit less from union coverage, whereas workers with less market power benefit more (Gomez & Lamb, 2019).

5.4 Answering our main research questions

The primary focus of this article is the potential compensating wage premiums associated with novel workplace risk attributed to COVID-19 and the role of unions in potentially moderating this relationship. In this regard, we find patterns consistent with previous results on workplace hazards and unions, and more importantly, with the predicted effects anticipated in our review of the relevant literature. We present our results in the order of main research questions identified in Section 2.3.

5.4.1 Did workers receive a pay premium during the COVID-19 lockdown?

The estimates in Table 3 (column 3) support the notion of a compensating wage associated with working during the COVID-19 pandemic. We find an overall earnings premium of roughly 5.3% for the variable Covid period, which covers all workers employed during the first 6 months of the COVID pandemic (March–September 2020) compared to the same period a year earlier. It bears mention that our results do not capture important selection effects, whereby particularly low earners may have been more likely to leave the labour force and/or become unemployed during the pandemic. Notwithstanding, the use of inflation-adjusted wage data shows significant year-over-year wage gains.

Interestingly, the interaction Covid period × Union is small but negative, and statistically significant at –2.7% (Table 3, column 3, row 6). This finding suggests that during the pandemic, unionized workers, on average, received a smaller wage increase than observationally similar non-unionized employees. This finding appears counterintuitive but is not for two reasons: First, collective agreements governing earnings in unionized workplaces are (re)negotiated upon expiration and therefore ad hoc wage changes are more difficult to implement. Second, we may reasonably conjecture that given the job security and other non-pecuniary provisions associated with unionization (i.e. greater adherence to health and safety), it is plausible that unionized workers would be
relatively less likely to leave the labour force or otherwise become unemployed during the pandemic, thereby requiring less of a wage inducement to remain.

Table 4 looks at estimated wage-risk trade-offs separately for male and female workers. In Table 4 (columns 1 and 2), we see this overall estimate of 5.3% is split at roughly 4.9% and 5.9% for males and females, respectively (Table 4, columns 1 and 2).

Looking across the earnings distribution in Table 5 (row 2), we find positive COVID-19 time-period wage premiums across the entire earnings distribution, with the largest at the bottom of the distribution. Workers at the 10th and 25th earnings percentiles each experienced a COVID-19 earnings advantage of roughly 7.6% and 6.9%, respectively. Moving up the earnings distribution, workers at the 50th and 75th percentiles enjoyed a roughly 5.3% and 4.3% COVID-19 earnings premiums, respectively. Finally, the highest earners at the 90th percentile experienced a 4.3% increase in earnings during the COVID-19 period.

This result likely reflects a combination of changes in the supply of employed workers during the pandemic as well as compensation policies that provided additional pay to frontline workers. Given the relatively low wages of this group, any increase in earnings resulting from employers’ offering additional pandemic pay would translate into a proportionally large increase in earnings relative to pre-COVID levels. The implementation of such programs was not uniform, and often for a very limited duration.

Notwithstanding the strong statistical association, any bump in earnings during the first 6 months of the COVID-19 pandemic must be interpreted with caution as it is very likely an artefact of changes to the labour supply, whereby particularly low wage earnings were more likely to leave the labour market or to become unemployed during the pandemic (see Lemieux et al., 2020). This can be seen qualitatively by looking at the side-by-side comparison of the hourly earnings distribution prior to and during COVID-19 shown in Figure 3. It is evident that mean earnings have risen primarily because the bottom of the earnings distribution has thinned considerably in the COVID-19 period. This reality has been noted by others looking at COVID-19 wage growth in the United States (Crust et al., 2020).

5.4.2 Do workers in high COVID-19 risk occupations earn more (less) than observably similar workers in lower COVID-19 risk occupations?

What about workers who are at greater risk of contracting COVID-19, do they earn a compensating wage premium for this newly observed form of hazardous work? For the whole sample and controlling for all relevant observable characteristics, workers in occupations at high risk of contracting COVID-19 (Covid risk high) earn roughly −7.6% less than workers in lower risk occupations (column 3, Table 3).

The wage penalty associated with high-risk work is dramatically larger for males than females at roughly −18.6% and −2.9%, respectively (columns 1 and 2 in Table 4).

In Table 5 (row 3), looking across the earnings distribution, workers at high risk of contacting COVID-19 experienced the largest earnings disadvantage – 11.4% and 11.2%, respectively – at the 25th and 50th percentiles. High-risk workers at the bottom of the distribution earned roughly 5.0% less than comparably lower risk workers, whereas the earnings penalties at the highest ends of the distribution were statistically significant, but quantitatively small, at 2.8% and 3.9% for those at the 75th and 90th percentiles, respectively (see Figure 4 bottom dashed line).

These negative CDs for workplace risk – although not predicted by standard theories based on competitive labour markets – are a frequently observed empirical result (Seibert & Wei, 1994).
### TABLE 4  Estimated log hourly earnings equations by sex

| DV: Ln hourly wages | Ordinary least squares |   |   |
|---------------------|------------------------|---|---|
|                     | Male                   | Female |   |   |
| Key independent variables | [1]                   | [2] |
| [Non-union] Union | 0.024**       | −0.038** |
|                     | (7.43)                  | (11.18) |
| [Pre-Covid period] Covid period | 0.049**       | 0.059** |
|                     | (19.04)                  | (20.97) |
| [Covid risk low] Covid risk high | −0.186**       | −0.029** |
|                     | (24.25)                  | (6.43) |
| Covid risk high × Union | 0.140**       | 0.101** |
|                     | (13.75)                  | (17.20) |
| Covid risk high × Covid period | 0.041**       | −0.025** |
|                     | (3.37)                   | (3.47) |
| Covid period × Union | −0.021**     | −0.034** |
|                     | (4.93)                   | (8.11) |
| Covid risk high × Covid period × Union | −0.048**     | 0.047** |
|                     | (3.06)                   | (5.33) |
| [Age 20–29] Age 30–59 | 0.141**       | 0.107** |
|                     | (47.71)                  | (38.22) |
| Age 60–64 | 0.066**       | 0.046** |
|                     | (14.01)                  | (9.85) |
| [Non-immigrant] Immigrant | −0.14**       | −0.141** |
|                     | (49.20)                  | (52.51) |
| [Full-time job] Part-time job | −0.247**       | −0.142** |
|                     | (51.25)                  | (52.38) |
| Demographic and human capital | Yes     | Yes |
| Job/workplace controls | Yes       | Yes |
| Industry controls | No         | No |
| Constant | 2.776**           | 2.581** |
|                     | (531.31)                 | (434.71) |
| R-squared | 0.33            | 0.41 |
| N | 273,307         | 263,017 |

Notes: The outcome variable is the natural logarithm of hourly earnings adjusted for inflation. *t*-statistics are in parentheses. Control variables included in the models, but not shown here, are those used on model 3, outlined in column 3 of Table 4. Models are weighted using STATA’s probability weights (p-weights) and the LFS sampling weight (FINALWT) rescaled for 14 months of data. Sample sizes represent the unweighted count of respondents in the dataset. Additional covariates not shown here include marital status, presence of children, region of residence, living in a large urban centre, survey month, tenure at current job (years) and firm size.

*p < 0.05.

**p < 0.01.
### Table 5: Unconditional quantile (RIF) regressions on the natural logarithm of hourly earnings

| Non-union | 10th  | 25th  | 50th  | 75th  | 90th  |
|-----------|-------|-------|-------|-------|-------|
| Union     | 0.102** | 0.105** | 0.038** | −0.07** | −0.165** |
|           | (37.31) | (31.86) | (10.21) | (16.83) | (38.46) |
| Pre-Covid period | 0.076** | 0.069** | 0.053** | 0.042** | 0.043** |
|           | (30.97) | (24.92) | (19.22) | (14.46) | (12.66) |
| Covid risk low | −0.05** | −0.114** | −0.112** | −0.028** | −0.039** |
|           | (7.16)  | (16.55) | (19.79) | (5.45)  | (7.59)  |
| Covid risk high × Union | 0.195** | 0.176** | 0.113** | 0.084** | 0.006  |
|           | (12.28) | (21.34) | (13.83) | (10.01) | (0.78)  |
| Covid risk high × Covid period | 0.041** | −0.025*  | −0.024** | −0.006  | −0.026** |
|           | (4.17)  | (2.32)   | (4.84)   | (1.04)  | (4.47)  |
| Covid period × Union | −0.058** | −0.04**  | −0.023** | −0.012  | 0.005   |
|           | (17.69) | (9.67)   | (2.65)   | (1.47)  | (0.57)  |
| Covid risk high × Covid period × Union | −0.051** | 0.013*  | 0.047**  | 0.052** | 0.014   |
|           | (4.69)  | (1.03)   | (3.80)   | (3.97)  | (1.09)  |
| Male      | −0.083** | −0.148** | −0.196** | −0.18** | −0.147** |
|           | (43.64) | (68.53) | (88.61) | (75.05) | (54.73) |
| Age 20–29 | 0.071** | 0.129** | 0.145** | 0.153** | 0.106** |
|           | (24.20) | (37.94) | (43.72) | (48.32) | (36.20) |
| Age 30–59 | 0.031** | 0.046** | 0.046** | 0.087** | 0.063** |
|           | (6.86)  | (9.04)   | (9.13)   | (16.98) | (12.06) |
| Age 60–64 | −0.082** | −0.142** | −0.176** | −0.188** | −0.121** |
|           | (33.79) | (50.79) | (60.45) | (49.99) | (34.72) |
| Non-immigrant | Immigrant | −0.082** | −0.142** | −0.176** | −0.188** | −0.121** |
|           | (33.79) | (50.79) | (60.45) | (49.99) | (34.72) |
| Full-time | −0.293** | −0.315** | −0.21**  | −0.076** | −0.013** |
|           | (65.44) | (79.60) | (63.84) | (23.99) | (4.09)  |
| Demographic and human capital controls | Yes | Yes | Yes | Yes | Yes |
| Job/workplace controls | Yes | Yes | Yes | Yes | Yes |
| Industry controls | No | No | No | No | No |
| Constant  | 2.344** | 2.359** | 2.62**  | 3.056** | 3.497** |
|           | (357.62) | (347.21) | (435.30) | (540.41) | (639.11) |
| R-squared | 0.15 | 0.25 | 0.28 | 0.23 | 0.13 |
| Observations | 536,324 | 536,324 | 536,324 | 536,324 | 536,324 |

Notes: The outcome variable is the natural logarithm of hourly earnings adjusted for inflation. t-statistics are in parentheses. Control variables included in the models, but not shown here, are those used on model 3, outlined in column 3 of Table 3. Models are weighted using STATA’s analytic weights (a-weights) and the LFS sampling weight (FINALWT) rescaled for 14 months of data. Sample sizes represent the unweighted count of respondents in the dataset.

*p < 0.05.  
**p < 0.01.
Indeed, CD theories built on imperfect markets and institutional differences between union and non-union sectors predict negative returns for non-union workers exposed to more hazardous work (Fairris, 1989, 1992). The differential for workplace risk turns out be very different for union and non-union workers, something we turn to in the next subsection.
5.4.3 Does union coverage (positively) moderate the pandemic pay premium for workers in occupations facing high COVID-19 exposure risk, as past research would predict?

The central hypothesis of this article is that because unionized workers have been found to receive a larger compensating wage differential for workplace risks, this should also be true for a ‘novel’ risk like COVID-19, since the factors that are related to COVID exposure are similar to those of other infectious diseases, such as an influenza, H1N1 and so on. Indeed, one of the mechanisms reviewed in relation to the union CDs for hazardous work was that unions are better at uncovering (and exposing to their members and management) risks that are otherwise unseen (or latent) at the workplace. If this is true, there will be greater demand from union members to address latent workplace risks (through wages and/or better safeguards) prior to an actual outbreak than among otherwise similar non-union workers. As a result, unions will be more likely to generate a larger CD for COVID-19 hazards.

To consider whether unions offered differential earnings benefits to high- versus low-risk workers, we interact collective agreement coverage (union) with our high-risk occupation dummy. Consistent with expectations, the Covid risk high × Union interaction is positive and significant (column 3, Table 3). This means that not only are high-risk workers with a union better off than high-risk workers without a union but, controlling for other observable characteristics, they also out-earn non-union workers in lower risk occupations.

Union membership results in a net increase in wages for unionized workers in occupations at high risk of being exposed to infection, which we interpret as ‘compensation’ for greater exposure to infectious diseases like COVID-19. In Table 3, column 3 results, the wage premium is roughly 10.3% higher for union members than it is for non-union members in the absence of a high COVID-19 exposure risk. Note that in model 3, this constitutes (more than) the entirety of the ‘main effect’ of unions (i.e. the traditional union wage differential). This is a significant interaction that positively moderates the (negative) effect of high COVID-19 exposure at work.

Looking at the models estimated for males and females separately, Table 4 (columns 1 and 2), the same patterns emerge. Though both interaction coefficients are positive, among males, the positive wage effects associated with unionization and high-COVID-19 exposure risk are larger than for females (14% vs. 10.1%, respectively).

Across the earnings distribution, Table 5 (row 4), we find that high-risk workers across all percentiles experience a net wage advantage over their non-union, lower risk counterparts, with the most sizable advantages found at the 10th, 25th and 50th percentiles of the earnings distribution. This can be seen graphically in Figure 4 (top dotted line), where the largest union wage premia are associated with high-risk workers at or below the median earner.

This pattern is remarkably similar to the union ‘main’ effect (Table 5, row 1). Consistent with what has been robustly supported by extant research, in Table 5 (row 1), we find that the union wage premium is large and significant in the lowest wage quantiles but diminishes around the 50th percentile and becomes a penalty at the highest points in the earnings distribution. At the 10th and 25th percentiles, unionized workers earn a 10.2% and 10.5% hourly earnings premium, which falls to 3.8% at the 50 percentile and turns negative –7.0% and –16.5% at the 75th and 90th, respectively. This distributional effect associated with the union wage premium is also visualized in Figure 4 (the middle light dotted line).
5.4.4 Any evidence of ‘hero pay’ for frontline workers?

In the early months of the pandemic, numerous media reports highlighted the increased compensation (often with government support) being offered to frontline workers. We explore this idea of ‘hero pay’ by interacting the variables for the COVID-19 period and being in a high-risk occupation (\(\text{Covid risk high} \times \text{Covid period}\)) to observe if, in fact, occupations at higher risk of contracting COVID were differentially rewarded during the pandemic. Overall, in Table 3 (column 3, rows 5 and 6), we find no significant CD associated with working in a high-risk occupation during the initial 6 months of the COVID-19 pandemic.\(^20\)

In terms of any CDs associated with the COVID-19 period and being in a high-risk occupation by sex and income percentile; males experience a 4.5% earnings advantage, whereas females in high-risk occupations earned 2.5% less during the pandemic (Table 4, columns 1 and 2). At the bottom 10th percentile of the earnings distribution, there is some evidence of ‘hero pay’ in that high-risk workers earn an additional 4.1% during the COVID-19 period. But workers in high-risk occupations above the 25th percentile all experienced a negative CD during the COVID-19 period (Table 5, row 5).

It is important to emphasize that interaction terms need to be interpreted considering the direct (main) effects of both COVID-19 time period and being in a high-risk occupation. For males at the mean (column 2, Table 4), the negative relationship between being in high-risk occupation and earnings is so large that neither the positive effect of the COVID-19 period in Table 4, row 2 (4.9%) on earnings, nor the interaction with high-risk of COVID exposure (4.1%) are sufficiently large enough to mitigate this disadvantage. The net result is that males in high-risk occupations still earn less than their counterparts employed in lower risk work; however, the disparity is smaller during COVID-19 than it was prior to the pandemic.

Without considering the effects of unionization, which we have already discussed, the only net positive compensating wage differential observed for working in a high-risk occupation in our analysis is, therefore, found among workers at the lowest end (i.e. bottom 10th percentile) of the earnings distribution who, during COVID-19 (Table 5, row 5, column 1), did receive earnings that nearly match their lower risk counterparts.\(^21\)

5.4.5 How does the inclusion of a standard measure of workplace risk affect our results?

In Table 3, we measured log hourly earnings and covid exposure risk, but we did not measure overall workplace risk as measured by injury or mortality rates. We address this here by using a standard measure of workplace risk, namely occupational risk of lost-time injury, introducing it as an additional regressor in our preferred estimation model [3]. Following Taigi (2015), we constructed a measure of lost-time injury rates which combines publicly available data on lost-time injuries by occupational group from the Association of Workers Compensation Boards of

\(^{20}\)The coefficient on the three-way interaction (\(\text{Covid risk high} \times \text{Covid period} \times \text{Union}\)) is small, but positive and statistically significant at roughly 2.0%. This suggests that relative to non-union counterparts, being unionized during the Covid pandemic period and working in a high-risk occupation was associated with a small but significant compensating wage advantage.

\(^{21}\)This result may of course be in part influenced by selection effects whereby the lowest wage earners may have been unemployed or left the labour market during the onset of the pandemic as noted earlier.
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Canada (AWCBC, 2021) and the 2016 Canadian Census (Statistics Canada, 2016). We averaged the total number of reported lost-time injuries in each major occupational group for the years 2016–2019 inclusive and divided each by the number of people employed in each occupation group as reported in the 2016 Census. The resultant variable is the average rate of lost-time injury for each major occupational group, which we then mapped onto the occupational group codes available in the LFS. The results are presented in Online Table A3.

We find that the inclusion of a standard injury risk measure does not reduce and render less significant the estimated coefficient for our Covid risk high × Union interaction. If anything, it appears to boost the direct association between unions and hourly wages and strengthens the interaction Covid/Union relationship; meaning that our measure of covid exposure risk was not inadvertently capturing an overall measure of workplace risk in our prior estimates. Except for the high covid risk occupation in model 4, which is now smaller in magnitude and statistically insignificant and the three-way interaction between Covid high risk × union × Covid period, which loses some significance, the other key variables remain stable following the inclusion of our lost-time due to injury variable.

6 | CONCLUSION

While the pandemic and subsequent institutional responses are continuing to evolve, our study is unique in considering the relationship between unions and compensating wage differentials for new workplace hazards, such as COVID-19 exposure risk. We triangulated data from a ranking system that scored occupations in the US O*NET database for COVID-19 exposure risk with Canadian occupational data contained in the PUMF of the LFS. This allowed us to categorize occupations based on COVID-19 exposure risk scores. We find a long-tailed distribution of COVID-19 risk across occupations, with most jobs at the lower end of the risk spectrum (scores under 40) with relatively few occupations accounting for most of the COVID-19 exposure risk. Further, we find that workers who are already more vulnerable in the labour market (i.e. the young, women, part-timers and immigrants) are also more likely have high COVID-19 exposure risk.

Looking at the inter-relationships among working during the COVID-19 lockdown period, occupational exposure risk, union coverage and hourly earnings, we find that wages were 5.3% higher between March and September 2020 as compared to the same 2019 pre-COVID period (a sizeable wage gain given that year-over-year inflation was close to 2%). When we look at the association between high-COVID exposure risks and wages, we find sizeable earnings disadvantages both at the mean (–7.6%) and across the earnings distribution. When workers in occupations with high COVID-19 exposure risk are covered by a union, however, they enjoy a sizeable earnings premium (10.3% on average). Furthermore, we find the counterbalancing effects of unionization for high-risk workers to be largest at the lower to middle ends of the earnings distribution (i.e. the 10th, 25th and 50th percentile of unionized earners receives a 9.5%, 17.6% and 11.3% wage premium, respectively), whereas the 90th percentile worker does not receive a statistically significant wage premium.

Our findings underscore the role of unions in protecting the most vulnerable along this new COVID-related dimension of workplace precarity. The current pandemic has sparked renewed interest in OHS and the notion of compensating wage differentials as seen through the lens of an unprecedented form of risk – exposure to COVID-19. We anticipate that the response from employers and unions will continue to evolve, and the estimates presented herein, specifically,
the extent to which workers in high-risk occupations experience CDs, will be revised as workers demand improvements in their terms and conditions of employment.

Recently, for example, UNIFOR, a large Canadian union representing grocery workers (among others), has called on employers to reinstate the initial ‘hero pay’ first given to frontline workers at the start of the pandemic (Robertson, 2021). Indeed, based on our key outcome of interest (i.e. hourly earnings), unionization was associated with a substantial compensating pay premium for workers with a high COVID-exposure risk. Unions typically also provide additional benefits, such as job security, paid sick days, closer adherence to OHS rules and extended health coverage that may be of even greater value to workers during these times. Considering the range of benefits available in most collective agreements, this widens the vulnerability gap between high-risk workers covered by collective agreements and those who are not covered. In so far as unions negotiate improved terms and conditions of employment for their members, efforts to organize workers at high risk of COVID-19 are critical during the pandemic and beyond since, even prior to COVID-19 occupations at high risk of infectious-disease exposure were associated with statistically significant earnings penalties as compared to their observably equivalent lower risk counterparts. Non-unionized employers are of course able to provide CDs on their own and indeed, may have a relatively easier time adjusting compensation policies. In practice, however, resource constraints and competitive pressures may make such adjustments onerous, particularly for small businesses.

As we have seen, in the absence of legislative requirements and/or enforcement mechanisms, such ‘hero pay’ premiums are arbitrary and short-lived. Time will tell if market mechanisms adjust for the novel COVID-19 exposure risk and whether high-risk occupations will have difficulties attracting and retaining qualified personnel. In the interim, for those workers in high COVID-19 risk occupations without collective agreement coverage, statutory regulations adding provisions for paid sick days to employment standards legislation, along with government subsidies for small employers, would do well to not only protect the health and safety of workers and their families, but also to mitigate the spread of the virus.

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DATA AVAILABILITY STATEMENT
This research uses data from the Public Use Microdata Files of the Canadian Labour Force Survey (LFS). Data is owned by statistics Canada and made accessible to participating institutions through the Data Liberation Initiative. LFS data were combined with COVID-19 occupational risk scores computed by Visual Capitalist (Lu, 2020) and publicly accessible from: https://www.visualcapitalist.com/the-front-line-visualizing-the-occupations-with-the-highest-covid-19-risk/. Data on Workplace Injury and Fatality rates are publicly available from the Association of Workers Compensation Boards of Canada at https://awcbc.org/en/statistics/#nwisp. Occupational information from the Canadian Census is publicly available from https://www150.statcan.gc.ca/n1/en/catalogue/98-400-X2016295. All errors are exclusively our own.

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