# Gender Differences in Sorting on Wages and Workplace Safety*

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**Abstract**

We study gender disparities in labor market sorting on wages and workplace safety. Using Brazilian matched employer-employee data, we show that when women change jobs they move toward safety with little change in wages. Although safety has little direct wage effect, its outsized influence on job choices segregates women and men across establishments and contributes to large indirect effects on establishment wage premia, which explain nearly one-third of the gender wage gap in Brazil. Employment segregation becomes more pronounced over the career, as sorting associated with safety and establishment pay contribute to the expansion of gender wage disparities.

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# 1 Introduction

There is a growing consensus that observed differences in earnings between women and men have little to do with skill. Rather, these differences arise from the kinds of jobs women and men hold, and the way they are compensated in those jobs. Holding participation and labor force attachment constant, women in many countries tend to work in lower-paying occupations (Cortes and Pan 2018), industries (Goldin et al. 2017), and firms (Bayard et al. 2003; Card et al. 2016). It remains unclear whether gender differences in labor market sorting reflect the preferences of women and men for particular types of jobs, or if they reflect discriminatory constraints on employment opportunities. In particular, preferences for non-wage job amenities could generate differences in sorting, but such amenities are often unmeasured.

We revisit the role of job amenities in affecting differences in labor market outcomes between women and men, focusing on a specific amenity—workplace safety. Hersch (1998) and DeLeire and Levy (2004) argue that women have very different preferences for safety than men, and that these preferences drive substantial occupational sorting. We build on this literature using matched employer-employee data from Brazil to graphically summarize patterns of job transitions. We show that sorting on safety is highly systematic and salient, even relative to sorting on establishment wage premia, a topic that has been extensively studied in the labor economics literature. The stark gender difference in sorting on safety relative to wages, presumably caused by preferences or employer discrimination, segregates workers across firms, and lead women to be over-represented at low-wage establishments (Lang and Lehmann 2012). This contrast in establishment assignment explains nearly one-third of the gender wage gap.

Figure 1 motivates our analyses in the remainder of the paper. While it appears complicated, the figure illustrates two simple facts—women and men have very different joint distributions of risk and wages, and exhibit very different patterns of choices when moving between jobs.

To interpret the figure, the horizontal axis lists quintiles of fatality risk across all jobs, and the vertical axis lists quintiles of wages. For each fatality rate and wage quintile pair, the figure's interior plots vectors representing the average gradient (direction and relative magnitude) of changes in wages and fatality risk when workers make job-to-job transitions originating from the quintile pair.[1] To the left of the vertical axis, we show overlapping density plots of female and male wages, with similar risk density plots below the

[1]The gradients for transitions by female (male) workers are represented by black (blue) arrows. Gradient vector lengths are proportional to the magnitude of the change in fatality rates or wages, normalized by the corresponding standard deviation of changes in fatality rates or wages.

horizontal axis.

The gradient field shows that for nearly all origin points, when women change jobs they move very systematically toward the left, indicating that the destination job has a lower fatality rate, with little change in the wage dimension. In contrast, male job transitions are far more consistent with mean reversion—risk increases when workers originate at safe jobs, and decreases when they originate at risky jobs. Moreover, men in low-wage jobs move on average several times further in the vertical direction, while women experience only modest wage increases. When originating in safe jobs, women's jobs have very little change in characteristics, regardless of the wage. The marginal wage distribution to the left is consistent with an 18 percentage point gender wage gap in Brazil, while the distribution below the figure indicates that women are employed in much safer jobs than men.

These mobility patterns suggest women consider very different sets of jobs and firms than men, either by choice or by exclusion. As we will show in Figure 2, workplace safety plays a role in this sorting process that is as systematic and salient as patterns of job mobility based on establishment wage premia. In jobs (industry-occupation pairs) in which women are safer than men, there is substantially greater gender segregation across establishments. A consequence of this segregation is that women end up employed in firms that on average pay lower wages to all workers, both female and male. We find that this differential sorting across establishments explains 32 percent of the total gender wage gap in Brazil, about twice the share found by Card, Cardoso and Kline (2016) in Portugal.[2]

A potential explanation for the different sorting patterns in Figure 1 is that women are not compensated for accepting risk at the same rate as men. Although gender differences in compensating wage differentials are a consensus empirical finding in the literature,[3] they are inconsistent with basic hedonic wage theory if women and men provide substitutable labor, and therefore share a common "marginal worker" whose preferences define the hedonic equilibrium (Rosen 1974). We rule out this explanation for the observed sorting patterns, and show that women and men earn equivalent rates of compensation per unit of risk in our setting. Although we can replicate the gender gap documented in the empirical literature, we show that it can be explained by two important estimation issues. First, previous estimates do not account for the non-random assignment of workers to jobs on the basis of unobserved worker, establishment, and occupational characteristics, or allow this assignment process to itself be gender-specific. Second, we improve upon the

[2]Within establishments in Brazil, we show that women and men earn roughly similar wages conditional on their occupation, measured, and unmeasured characteristics.

[3]See Hersch (1998) and Leeth and Ruser (2003).

measurement of safety by calculating precise gender by industry by occupation by time measures, which have not previously been available.[4] After correcting these estimation and measurement issues, the gender gap in compensating wage differentials is reduced to a precisely-estimated zero in our setting.

Our paper contributes to a small, but growing literature on the role of workplace sorting in the propagation of the gender wage gap. Card et al. (2016) discuss (and discard) the possibility that observed differences in firm-specific pay for women and men in Portugal reflect compensating differentials for hours requirements. Hotz et al. (2017) study the contribution of family-friendly policies to career outcomes for women immediately after childbirth. We are among the first to focus on worker sorting on the basis of a specific, observable, job disamenity—workplace safety—that has been shown in prior work to be highly salient in women's labor market decisions (Hersch 1998).

While our primary interest is in sorting across firms, our analysis also complements recent work documenting differences between men in women in sorting within firms. Survey evidence by Wiswall and Zafar (2018) suggests that women and men have very different preferences for workplace amenities, and that women are less inclined to sort into jobs based on earnings growth. Mas and Pallais (2017) use experimental variation to recover preferences for job amenities, and find important gender differences in the willingness-to-pay to avoid irregular work schedules. Goldin, Kerr, Olivetti and Barth (2017) show the gender gap is driven in part by differences in pay growth within firms, after controlling for sorting across firms. Merlino, Parrotta and Pozzoli (2014) find gender wage gaps may be driven by differences in the propensity for women to be promoted within the firm, while Bartolucci (2013) show women have lower bargaining power than men, but are also more mobile. Unlike these studies, we are uniquely able to consider comprehensive patterns of labor market sorting, and to estimate the importance of sorting on job safety relative to wages.

## 2 Data

We use matched employer-employee data from Brazil's *Relação Anual de Informações Sociais* (RAIS) from 2003—2010. The RAIS is an administrative census of every formal-sector job, conducted annually by the Brazilian Ministry of Labor and Employment for the purpose of administering tax and transfer programs. We use these data to measure workplace sorting, and to decompose the gender wage gap. We also are able

[4] We follow Tsai, Liu and Hammitt (2011) and Lalive (2003), who showed that using overly coarse industry-occupation summaries of fatality risk generate substantial aggregation bias when estimating compensating wage differentials.

to use the reported cause of job separation in RAIS to construct highly detailed, gender-specific, measures of fatality risk.

Starting with the raw RAIS data, we construct a worker-year panel focused on individuals aged 23-65 who are employed for at least 1000 hours per year in at least one full-time job, defined as 30 or more contracted hours per week.[5] When a worker holds multiple jobs in the same year, we restrict attention to the job with the highest estimated annual earnings.[6] We exclude government jobs, and temporary employment contacts. As we describe below, to reduce error in measuring fatality rates, we also exclude jobs in 2-digit industry by 3-digit occupation cells that contain fewer than 10,000 full-time full-year-equivalent workers of either gender.

In our analysis, we estimate establishment-specific wage effects separately for each gender. This is only possible in establishments where at least one worker of each gender either leaves to, or arrives from, another establishment. In the language of Abowd et al. (2002), we focus on establishments that are in the largest connected component of the realized mobility networks generated by male job transitions and by female job transitions. For women and men, taken separately, the largest connected components contain 91% of observations in the attached dominant jobs sample. The intersection of these sets, the dual connected set, contains 8.2 million observations for women, or 91% of the largest female connected set, and 14.6 million observations for men, 72% of the largest male connected set.

We report summary statistics in Appendix Table A1 by gender for the full population, the subset of dominant jobs and attached workers, and for the dual connected set of stayers and movers. Jobs in the dual connected set are higher-paying and have longer tenure than dominant jobs not in the dual connected set. Relative to the population, women in our analysis sample are slightly older, more likely to be white, less educated, have more experience and job tenure, higher hourly wages, and are more likely to be in jobs with zero observed fatalities. These differences are due primarily to our restriction to larger industry-occupation cells.

[5]Blau and Kahn (2016) show that part of the gender wage gap can be attributed to a persistent gender division of household labor, and resulting differences in labor force attachment. Our sample restrictions reduce the influence of gender division of labor, but focus on the subset of women with strong attachment.

[6]This approach follows Abowd, Kramarz and Margolis (1999), Woodcock (2008), and Card, Heining and Kline (2013). The data only report average monthly earnings, so the estimated annual earnings are equal to monthly earnings times the number of months employed.

## 3 Gender Differences in Workplace Safety

The RAIS records whether each job ends due to a fatal workplace-related accident, and thus serves as a census of individual fatal occupational injuries. We use this information to construct gender-specific fatality rates in each year for 11,440 two-digit industry by three-digit occupation cells. Since fatalities are uncommon, we follow the literature in combining annual data to construct three-year moving averages of the number of fatalities per 100,000 full-time full-year-equivalent workers. In contrast, prior research has generally been limited to industry- or occupation-specific risk measures that do not differentiate by gender.[7]

Our improved measurement of fatality risk reveals large differences in job safety between women and men. Within the same industry-occupation cell, women have 38% lower fatality rates than men on average. This is potentially consistent with women being more cautious than men, or that women and men perform different tasks even within narrowly-defined jobs. We revisit this issue in Section 7.

In addition to being safer within industry-occupation cells, women also sort more strongly across cells on the basis of safety. This is consistent with well-documented gender differences in risk preferences generally (Blau and Kahn 2016; Bertrand 2011). Using fatality rates measured by pooling data from both genders, the average fatality rate for men is about three times higher than that of women. Using gender-pooled fatality rates, the measured disparity is even stronger: men are employed in jobs with average fatality rates about 8 times higher than women. Additionally, women are three times more likely than men to be employed in jobs with a zero measured fatality rate.[8] We also estimate that doubling the female fatality rate is associated with a 15.0% reduction in the female employment share (Appendix Figure A2a,) and that the growth of female employment over time is disproportionately concentrated in jobs that are safer (Appendix Figure A3.)

## 4 Wage Decomposition Model

We follow the empirical estimation approach proposed in Lavetti and Schmutte (2018), which we refer to as the *orthogonal match effects* (OME) estimator. This two-step specification allows time-varying characteristics $x_{it}$ to be arbitrarily correlated with unobserved worker, establishment, and job-match effects. It also allows fatality rates to be correlated with latent worker and establishment wage effects. Controlling for establishment heterogeneity in wages matters because higher-paying employers tend to offer safer jobs,

[7]One notable exception is (Hersch 1998), which uses gender-specific measures of injury risk within coarse occupation categories, but does not disaggregate by industry.

[8]See Appendix Table A1 for exact figures and more details.

and workers frequently move toward jobs with better wages and better amenities. Importantly for our application, this specification also controls for all establishment-level job amenities (such as the family-friendly policies studied by Hotz et al. (2017)) mitigating a common source of omitted variable bias in estimates of compensating differentials.

In a first step of the model, we remove the effect of time-varying observables using within-match variation by estimating:

$$w_{it} = x_{it}\beta^{g} + \tilde{\gamma}^{g} a_{c(i,t),t} + \Phi_{i,Jk(i,t)} + \varepsilon_{it} \tag{1}$$

where the dependent variable, $w_{it}$, is the log hourly wage of worker $i$ in year $t$. The vector of control variables, $x_{it}$, includes a set of indicators for each year of labor market experience (censored at 30 years) fully interacted with gender, and year effects. The coefficient vector $\beta^{g}$ denotes that each coefficient in $\beta$ is free to vary by gender. $a_{c(i,t),t}$ is our gender-specific measure of the risk of fatal injury when the worker is employed in year $t$ in the industry-occupation cell given by the function $c(i,t)$ which maps panel indices to cells. $\Phi_{i,Jk(i,t)}$ denotes the match effect between worker $i$ and the $Jk$ establishment-occupation pair at which worker $i$ is employed in period $t$.

In the second stage, we estimate a decomposition of the components of variation in wages across jobs, while conditioning on the unobserved characteristics of establishments that may affect job mobility choices:

$$w_{it} - x_{it}\widehat{\beta^{g}} = \pi^{g} z_{it} + \gamma a_{c(i,t),t} + \gamma^{f} a_{c(i,t),t} * Female_{i} + \theta_{i} + \psi^{g}_{J(i,t)} + \varepsilon_{it} \tag{2}$$

This second stage is the two-way fixed effects decomposition popularized by Abowd et al. (1999). Variation in this model that contributes to the identification of $\gamma$ comes from job changes in which a worker moves across industry-occupation cells, which could occur either within or across establishments. $z_{it}$ includes gender-by-year effects and one-digit occupation code effects interacted with gender, to account for broad differences in job characteristics other than safety. In this model, the unobserved error depends on a component that is common to all observations for the same worker, $\theta_{i}$, and a common establishment wage effect $\psi^{g}_{J(i,t)}$ that can vary arbitrarily by gender. The index function $J(i,t)$ maps panel observations of the job held by worker $i$ in year $t$ to establishment identifiers.

Lavetti and Schmutte (2018) develop a theoretical model characterized by differentiated employers, frictional search, and endogenous amenity choices, and discuss the relationship between the structural equilibrium log wage equation in the model and the OME wage model.

**Key Assumptions** The key identifying assumptions are that $\mathbb{E}\left[\psi^{g}_{J(i,t)}\varepsilon_{it}\right]=0$ and $\mathbb{E}\left[a_{c(i,t),t}\varepsilon_{it}\right]=0$. A sufficient condition for satisfying the first moment restriction is that job mobility across establishments is independent of the wage residual. We replicate the event study diagnostics used by Card et al. (2013) to evaluate comparable exogeneity conditions, and find no evidence of violations of exogenous mobility (Appendix Figure A4). In other work (Lavetti and Schmutte 2018), we exhaustively assess empirical evidence that generally supports both exogeneity conditions using data on male workers in Brazil. The assumptions regarding establishment effects have also been thoroughly investigated using RAIS data by Alvarez, Benguria, Engbom and Moser (2018) and Gerard, Lagos, Severnini and Card (2018).

The second restriction is satisfied if the wage equation is correctly specified with respect to the hedonic price of risk. Lavetti and Schmutte (2018) investigate the plausibility of this condition using a network-based IV model, and find that instrumenting for $a_{c(i,t),t}$ has no effect on estimates, consistent with this restriction holding.

A separate estimation concern is whether there are sufficient job-to-job moves to obtain consistent estimates of establishment effects. The potential for limited mobility bias is relevant for our comparison of the distributions of estimated establishment effects for women and men (Andrews et al. 2008). To correct for limited mobility bias in $\Psi^{g}_{J(i,t)}$, we implement the split-panel jackknife method proposed by Dhaene and Jochmans (2015) using sample periods 2005-2007 and 2008-2010, normalizing $\Psi$ to be mean-zero in each sample.

Finally, we also have to ensure that differences in the establishment effects, $\widehat{\Psi^{g}_{J(i,t)}}$, estimated for men and for women have a valid interpretation. Intuitively, establishment effects are identified by comparing the wages of workers who move between different establishments. Since we (almost) never observe a single worker transitioning from receiving a female establishment effect to receiving a male establishment effect, a direct comparison of female and male establishment effects is not possible.[9] Under the standard normalization (Abowd et al. 2002), the gender-specific establishment effects have mean of zero within each gender; comparing establishment effects across genders requires a normalizing assumption.

Following Card et al. (2016), we normalize the gender-specific establishment effects by assuming that for industries with the lowest average values of $\Psi^{g}_{J(i,t)}$ there is no rent-sharing between establishments and workers, so that any relative differences in establishment effects between women and men in these industries

[9] Notwithstanding infrequent within-worker changes in reported gender documented by Cornwell et al. (2016).

can be interpreted as level differences, providing a common normalization for the two disconnected sets. We implement this normalization by calculating the average $\Psi^g_{J(i,t)}$ by two-digit industry, and estimating the average difference between female and male $\Psi^g_{J(i,t)}$ in the five industries with the lowest establishment effects.[10]

# 5 Simultaneous Sorting on Wages and Fatality Risk

We provide a simple framework to characterize job mobility patterns of workers moving across a two-dimensional attribute space, and then show empirical estimates of sorting on establishment pay and safety.

Workers may have different relative preferences for wages and for safety. In particular female and male workers may differ on average in their preferences. On the employer's side, jobs are differentiated by $\psi$, the average log wage premium associated with the establishment, and $a$, which measures the bundle of job attributes correlated with fatality rates. Jobs are characterized by the bundle $\mathbf{y} = (\psi, a)$. Workers are characterized by $\mathbf{x} = (\theta, g)$, which measures their skills, $\theta$, and gender, $g$. A match between a worker with attributes $\mathbf{x}$ and job with attributes $\mathbf{y}$ generates surplus $\sigma(\mathbf{x}, \mathbf{y})$. The job surplus, $\sigma$, can vary across genders due to differences in relative preferences for safety, productivity differences between women and men, or due to employer discrimination.[11]

As in the multidimensional sorting model posed by Lindenlaub and Postel-Vinay (2017), employed workers switch between jobs when the surplus at the destination job exceeds that at the origin job: $\sigma(\mathbf{x}, \mathbf{y^d}) > \sigma(\mathbf{x}, \mathbf{y^o})$. This process generates conditional transition probabilities $Pr[\mathbf{y}|\mathbf{y^o}, \mathbf{x}]$ that describe the probability a worker moves to a job with bundle $\mathbf{y}$ given he or she has attribute $\mathbf{x}$ and is currently employed in job $\mathbf{y^o}$.

Without additional structure, we can characterize sorting by the vectors $\{\mathbb{E}[\Delta\psi|\mathbf{y^o}, \mathbf{x}], \mathbb{E}[\Delta a|\mathbf{y^o}, \mathbf{x}]\}$, which describe the average change between the destination and origin jobs in wages and fatality rates, conditional on worker characteristics. We empirically estimate these vectors to characterize the full distribution of revealed job change behavior in the $\psi$ and $a$ dimensions as a function of gender, $\theta$, and the characteristics of origin jobs.

Figures 2a to 2d each display gradient fields describing the average change in establishment effects

[10]These industries include, from lowest to highest, petroleum coke production (50% women); sports, recreation, and leisure activities (59% women); printing and reproduction (79% women); chemical manufacturing (23% women); and food, lodging, and hospitality (49% women). We also estimate the model using only the industry with the lowest establishment effect, and the estimates are very similar.

[11]For example, this representation allows, as suggested by Garen (1988), that some workers may be relatively more productive in dangerous employment than others, or, conversely, some workers are better at managing risk.

$\Psi^g_{J(i,t)}$, and in fatality rate, $a$, at starting points defined by the deciles of $\Psi^g_{J(i,t)}$ and $a$. We plot these separately for groups of women and men with estimated person effects ($\theta$) above and below the mean. To help visualize differences in the magnitudes of the gradient vectors, we plot the field of gradients over the level sets describing gradient magnitudes (gradient vector norms).

The most striking contrast between women and men is in movement toward safety. Women almost never move toward more dangerous jobs on average. Even low-wage women, employed in the safest jobs, and in the lowest-paying firms, do not accept meaningful increases in risk. By contrast, both low- and high-wage men tend to move toward more dangerous jobs, even those originating in high-wage firms. Overall, women's job movements are more strongly characterized by vertical movements, consistent with a job ladder model in which women seek employment at high wage firms but are less willing accept increases in risk to achieve this outcome.

The figures also show a striking contrast between the mobility of high-wage and low-wage workers that is common across genders. Job movements of high-wage women and men are strongly in the direction of higher-paying firms. By contrast, the job movements of low-wage women and men are generally stronger in the direction of safety. Low wage women and men are also both observed to move from higher-wage to lower-wage firms, which could either reflect assortative matching, or involuntary job transitions. However, we find similar sorting patterns when restricting the sample to transitions initiated by worker resignations, which are plausibly more likely to be voluntary (Appendix Figures A10 and A11). Consistent with the overall contrast in sorting between high-wage and low-wage workers, our estimates imply a positive correlation between $\theta$ and $\Psi^g_{J(i,t)}$.

Men move toward riskier jobs when originating below the 7th decile of the $a$ distribution, but the direction of sorting switches at the top of the distribution. The magnitude of female gradients (vector norms) are on average much smaller than those for men, implying that origin and destination jobs are more similar. The vector projection of the female gradient in the vertical dimension is also smaller on average, suggesting that women are less likely to move toward establishments with higher average pay.

To ensure that these patterns are not driven by unmeasured establishment-level heterogeneity in worker safety, we test whether establishment-occupation cells with a higher share of female employees than the industry-occupation average have lower fatality rates (for either men or women). We find no relationship between establishment-level fatality rates and gender shares.

One caveat to interpreting these figures is that they assume the presence of some form of labor market friction, otherwise workers would always choose jobs with the highest combination of amenities. However, if frictions differ by gender then mobility patterns are not necessarily indicative of productive complementarities or differences in average preferences. For example, gender-based hiring discrimination could alter gradients in the $\psi$ dimension. The figures do not provide information about underlying causes of sorting.

# 6 Gender Wage Gaps and Implicit Prices

One possible explanation for the differences in worker sorting in Figure 2 is that they could arise due to different incentives, rather than preferences. If women receive smaller wage gains than men when moving between the same pair of firms, or receive lower compensating differentials per unit of fatal risk, then differences in sorting could be straightforward responses to facing different implicit wage prices. We find that the data do not support this explanation.

## 6.1 The Role of Firms in the Gender Wage Gap

To quantify the contribution to the gender wage gap of establishment assignment, we begin by estimating an Oaxaca-Blinder decomposition of the OME model components. This decomposition shows that the assignment of women and men to different establishments explains 32 percent of the total gender wage gap in Brazil, while differences in establishment wage premia conditional on assignment explains only 4 percent.

That is, we find that a woman assigned to the same establishment as a man receives, on average, a similar, though slightly lower, establishment wage premium (See Appendix Figure A6). One exception is low wage establishments, where women earn higher wage premia than men. The net effect of within-establishment gender differences in establishment effects explains about 0.7 percentage points to the overall gender wage gap in Brazil. If women were paid the male establishment wage effect, holding fixed establishment assignment, their wages would *decrease* by 1.6 percentage point. Interestingly, the same pattern is true for men—if men were paid the female establishment wage effect, their wages would also decrease by 2.1 percentage points. That is, on average workers in Brazil are concentrated in jobs in which workers of the same gender earn modestly higher establishment wage premia than workers of the opposite gender.

Despite relatively small net gap within establishments (0.7 pp), sorting across establishments contributes substantially to the gender wage gap. The average female establishment wage premium is 21.6 log points while the average male establishment premium is 27.3 log points (Appendix Table A2). The male distribution of establishment effects is shifted roughly uniformly to the right relative to the female distribution (Appendix Figure A7a), with a mean difference of 5.7 log points, or 32 percent of the wage gap. Partially offsetting this, women in our sample have higher levels of education than men (the effect of which is absorbed by $\theta$), and hold different occupations, which narrows the wage gap by 2.6 percentage points.

Figure 3 summarizes how differences in establishment assignment contribute to the evolution of the gender wage gap over the career cycle, and by education. First, among college educated workers, the majority of the difference in establishment pay arises due to differential sorting after age 25, while about 3 percentage points comes from persistent early-career differences in assignment. Second, the gap between female and male establishment effects is more than twice as large among workers without a college degree. Among workers with lower educational attainment, a relatively large share of the difference in average establishment pay is already evident by age 25, suggesting that among these workers the establishment component of the earnings gap has more to do with factors that affect initial assignment, such as preferences, skills, or discrimination. Nonetheless, in an absolute sense the expansion of the establishment pay gap associated with sorting is even larger among less educated workers. Finally, it is notable that there is strong sorting of workers overall on the basis of educational attainment. Since women in Brazil are about twice as likely as men to have a college degree (18% versus 9%), this suggests that the education-adjusted gender gap in pay associated with assignment to establishments would be even larger than the unadjusted gap we report.

### 6.2 Gender Differences in the Implicit Price of Safety

We now ask whether women sort more strongly toward safety because they are paid less than men for accepting fatality risk. Table 1 reports estimates of the compensating wage differential, $\gamma$, from our preferred OME specification. We find that when we account for the large differences in fatality risk across genders, the estimated compensating differentials for men and for women are indistinguishable.

First, in columns (1)-(2) we present estimates separately for men and women using fatality risk measures that pool data from both genders. The coefficient estimate for men, 0.233, suggests that an increase in the average fatality rate of one death per 1,000 full-time equivalent worker-years is associated with an approximately 23 percent increase in the hourly wage.[12] The compensating differential for women, 0.161,

[12]This is estimate is comparable, though slightly larger, than the estimate of 0.17 from Lavetti and Schmutte (2018) due to

is 31% lower, consistent with qualitative patterns from prior literature.

In columns (3)-(4) we re-estimate the same specifications using our preferred gender-industry-occupation fatality rates. Column (5) presents estimates from the full combined specification, which includes both women and men, and allows $\gamma$ (in addition to establishment effects and all other control variables) to be gender-specific. Across all three specifications, the estimated compensating wage differentials for women and men are equal, at $\gamma = 0.174$. In Column (5), the estimated deviation from this base parameter for women is 0.001 (SE=0.005), confirming that there is no economically meaningful difference in the compensating wage differential across genders.

This result clarifies somewhat puzzling evidence from previous studies on this topic. Hedonic theory suggests that the equilibrium compensating wage differential should be determined by the preferences of the marginal worker. Since non-segmented labor markets share a common marginal worker, the substitutability of female and male labor generally suggests that large gender differences in compensating wage differentials for safety should be unlikely to occur, even if women and men have different preferences for safety on average. Our findings suggest that more precisely measuring occupational safety, and correcting for the effects of endogenous assignment of workers to establishments, can help explain this puzzle.

## 7 Sorting on Safety, Establishment Segregation, and Sorting on Wages

To what extent is the establishment assignment component of the gender wage gap associated with sorting on safety? Figure 4 addresses this question, showing that sorting on relative safety is strongly connected to the segregation of women and men into different establishments, and that this establishment segregation measure is much stronger in industry-occupation cells in which men earn substantially higher establishment wage effects, $\Psi$.

In the figure, the vertical axis measures segregation using the dissimilarity index across establishment-occupations within each industry-occupation cell. The dissimilarity index is defined as:

$$D = \frac{1}{2}\sum_{\ell=1}^{K}\left|\frac{f_\ell}{F} - \frac{m_\ell}{M}\right|$$

where $K$ is the number of establishments within the cell, $f_\ell$ is the number of females employed in the given occupation in establishment $\ell$, and $F$ is the number of females employed in the cell, with equivalent

restricting the sample to the dual connected set of establishments.

definitions for $m_\ell$ and $M$. The index can be interpreted as the share of workers of either gender who must be reallocated to make the share of female workers in each establishment-occupation identical to the overall female share in the industry-occupation cell. The index ranges from zero, in which case there is no excess segregation, to one, in which case there is total segregation.

In Figure 4a the axis measures the female fatality rate minus the male fatality rate (from left to right the female fatality rate increases relative to the male rate). The figure shows that when women have a larger relative safety advantage, there is stronger segregation of women and men across establishments relative to the overall average gender share in the industry-occupation.

Figure 4b then shows that this establishment segregation is correlated with the gender gap in the average establishment wage effect in an industry-occupation cell. If women sort into jobs based on relative safety advantages, this sorting is correlated with an increase in gender segregation across establishments that tends to lead to men being overrepresented at high wage establishments. This combination of figures depicts the channel through which sorting on occupational safety increases the gender wage gap by affecting establishment assignment.

Figure 4a also helps explain why women appear to be safer than men within the same industry and occupation. If this pattern could be explained by gender-based task-assignment within industry-occupations, one may have hypothesized that in more gender-segregated establishment-occupation cells there would be less potential for such task-shifting, leading to smaller gender differences in safety. Figure 4a shows the opposite pattern. In the absence of task-specific data, we cannot definitively rule out this possibility in favor of alternative explanations like unmeasured sorting, or women behaving differently than men when performing similar tasks.[13] However, the figure suggests that whatever makes women safer is associated with gender-based sorting across establishments within narrowly defined types of jobs.

# 8 Discussion

A large literature in labor economics has documented the importance of sorting into high-wage and low-wage firms in explaining career wage dynamics and changes in earnings inequality over time. Our analyses reveal that for many workers, sorting on safety plays a strikingly large role in understanding job dynamics, and is roughly as salient in explaining mobility patterns as sorting on firm-level compensation. Although

[13]Cortes and Pan (2018), for example, use O*NET task data and show that women sort into occupations differently than men in part on the basis of the set of tasks associated with an occupation.

safety has a relatively small direct impact on wages, its outsized effect on sorting leads to large indirect increases in gender wage inequality, explaining nearly one-third of the entire gender wage gap in Brazil.

Of course, there are potentially many unobserved amenities that are correlated with safety and vary within establishments and occupation groups, and our estimates identify the combined effect of these amenities. Relatedly, Sorkin (2018) uses a revealed preference approach based on job mobility patterns, and infers that the net value of a bundle of unobserved establishment-level amenities explains over half the firm component of earnings variation. Taber and Vejlin (2016) estimate that one-third of job choices would have been different if workers cared only about wages. Our finding corroborate and extend this evidence to show that a specific observed amenity, safety, is, at a minimum, a surprisingly strong proxy for the set of amenities that explain workers' choices between jobs. Moreover, the strong degree of safety-related mobility appears to be of first-order importance for understanding how women and men sort differently through the labor market, and why their wages diverge when transitioning jobs.

Our findings are suggestive that policies aimed at improving workplace safety may alter the patterns of sorting across establishments, potentially reducing gender segregation in labor markets and the wage gap. A caveat to this possible implication is that, despite a 30% decline in the overall occupational fatality rate for women in Brazil between 2005-2010, the gender wage gap remained virtually unchanged. These safety improvements were one-sided, however, as the safety gap between female and male workers actually increased over this period. This leaves open the possibility that broad-based improvements in occupational safety may still have an equalizing force on labor market earnings potential for workers with different safety preferences.

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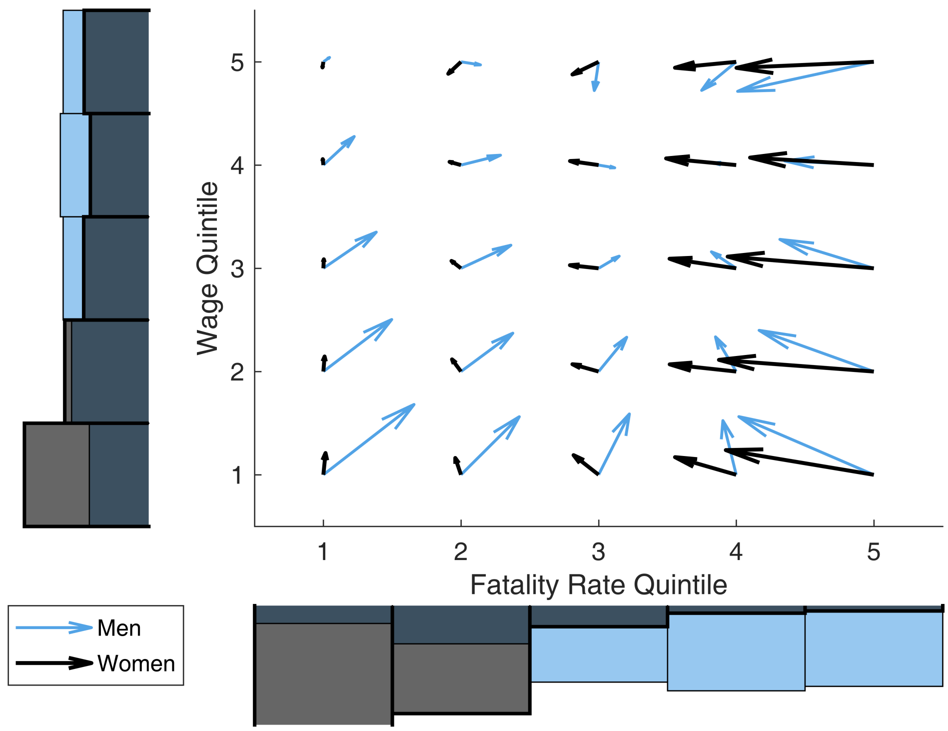
Figure 1: Job-to-Job Transition Gradient Field



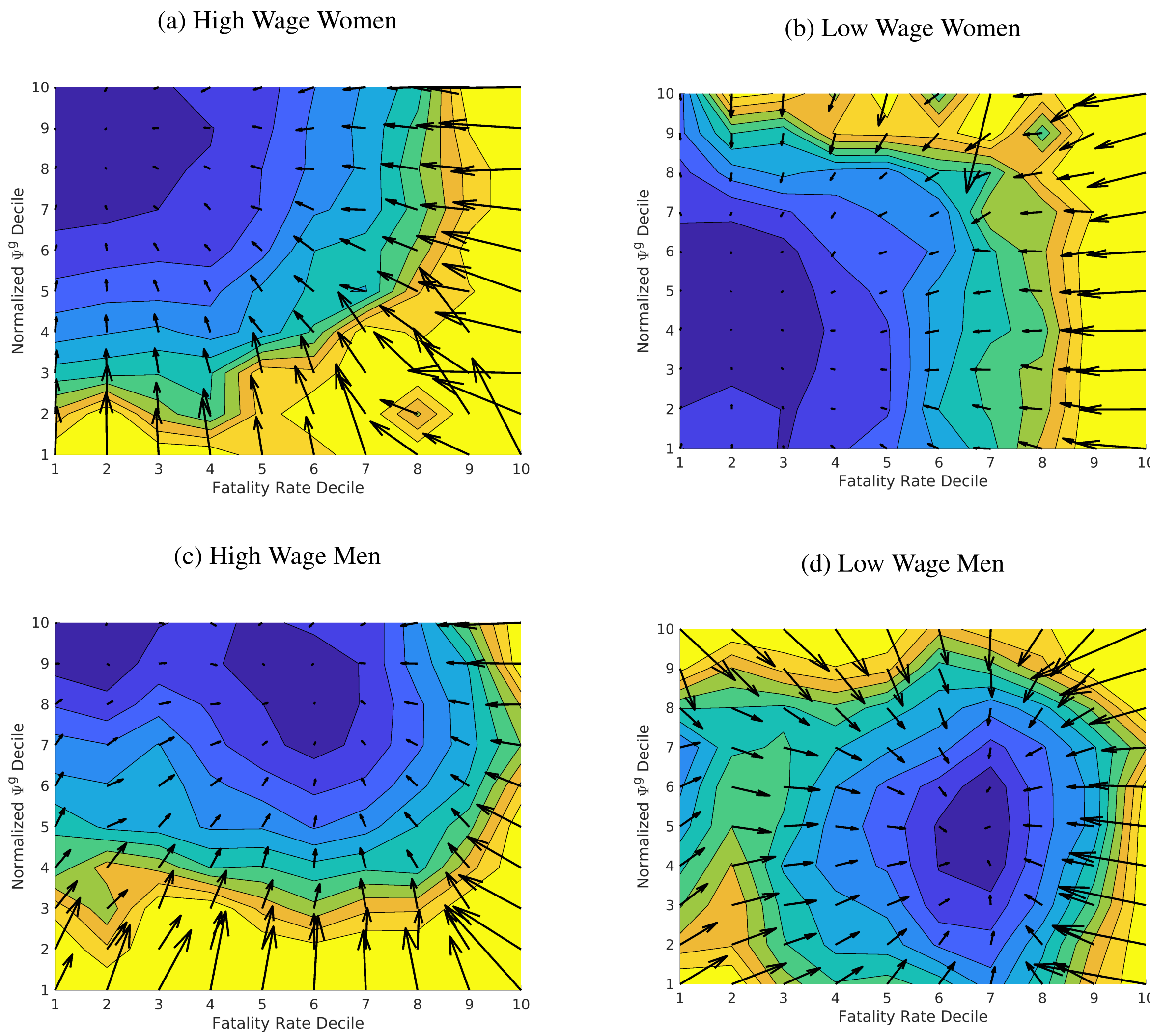

Notes: Sample includes attached full-time women and men between ages 23–59 employed in the largest dual connected set of establishments, and excludes origin jobs with zero fatality rate. Quintiles defined based on combined female and male distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Marginal density functions depicted at origin jobs.

Table 1: Gender-Specific Compensating Wage Differentials, OME Model

| | Fatality Rate Industry*Occupation | | Fatality Rate Gender*Industry*Occupation | | |
|---|---|---|---|---|---|
| | (1) Men | (2) Women | (3) Men | (4) Women | (5) Both |
| Fatality Rate | 0.233* | 0.161* | 0.174* | 0.174* | 0.174* |
| | (0.002) | (0.005) | (0.002) | (0.005) | (0.002) |
| Fatality Rate*Female | | | | | 0.001 |
| | | | | | (0.005) |
| VSL (million reais) | 3.41 | 2.06 | 2.55 | 2.23 | 2.43 |
| | [3.34, 3.47] | [1.94, 2.18] | [2.49, 2.60] | [2.11, 2.35] | [2.34, 2.53] |
| N | 13,985,793 | 8,131,646 | 13,985,793 | 8,131,646 | 22,117,439 |
| R-Sq | 0.959 | 0.970 | 0.959 | 0.970 | 0.971 |

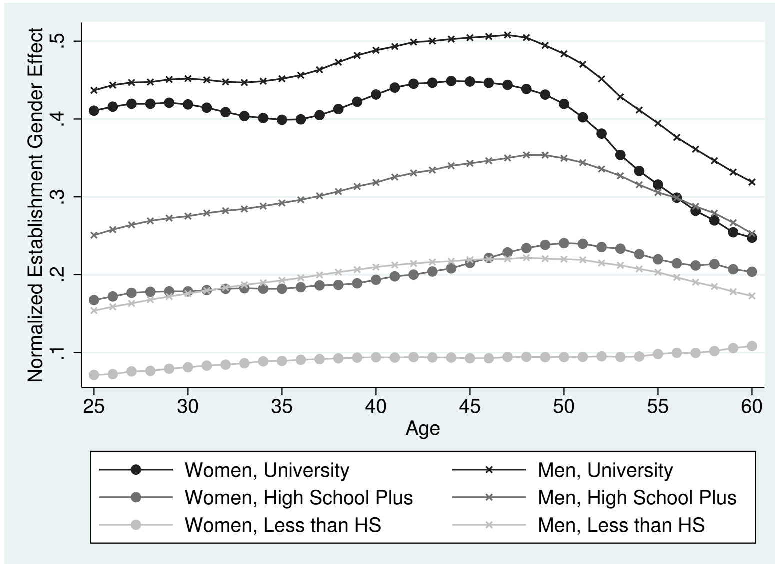
Notes: Analysis sample includes dominant jobs of attached workers in the dual connected set. Column 5 includes both women and men, with gender-specific establishment effects and gender-specific coefficients on all other control variables (year, experience, and occupation). 'Fatality Rate' is measured in deaths per 1,000 FTFY workers and is Winsorized at the 99th percentile. Log wages are Winsorized at the 1st and 99th percentiles. Implied values of statistical life (VSLs) are measured in millions of 2003 reais and calculated at the mean hourly wage in the corresponding sample, with 95% confidence intervals reported in brackets. VSLs are computed by converting the estimated semi-elasticity into a marginal effect $\frac{\partial w}{\partial a}$, and rescaling. Specifically, $\widehat{VSL} = \overline{w}\widehat{\gamma} * 2,000,000$, where the scaling factor, 2,000,000 is the product of 2,000 hours per FTFY worker with 1,000 workers, since $\gamma$ is the wage effect of an increase in number of fatalities per 1,000 workers. * Indicates significance at the 0.01 level.

Figure 2: Job-to-Job Transition Gradient Fields: Establishment Pay versus Risk
High Wage versus Low Wage Workers




Notes: Samples include attached full-time women/men between ages 23-59 employed in the largest dual connected set of establishments, and excludes origin jobs with zero fatality rate. Deciles defined based on gender-specific wage and fatality rate distributions. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

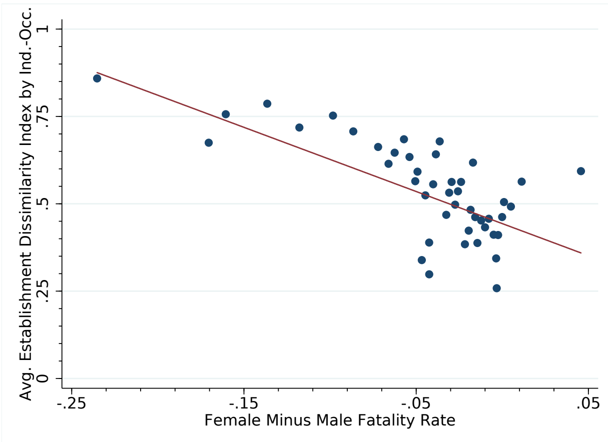
Figure 3: Average Establishment Wage Effects by Gender, Age, and Education



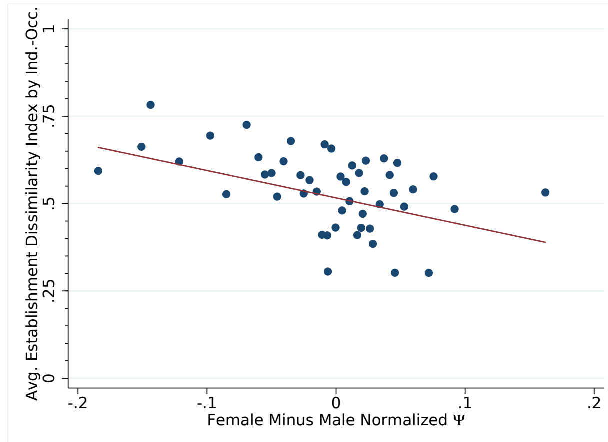

Notes: Figure depicts average normalized gender-specific establishment wage effects by age and education for women and men. Estimates based on the dual connected set.

Figure 4: Differences by Gender in Fatality Risk and Establishment Segregation

(a) Establishment Segregation and Safety




(b) Establishment Segregation and Normalized Ψ



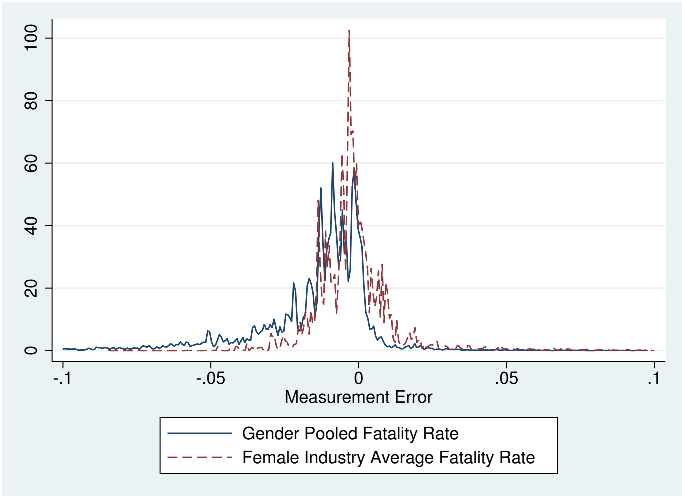

Notes: Vertical axis in both figures is the average establishment dissimilarity index for female and male workers within a 2-digit industry and 3-digit occupation. Horizontal axis in Figure 4a is the female minus male fatality rate, measured in deaths per 1,000 FTFY workers. Horizontal axis in Figure 4b is the female minus male average normalized Ψ in the corresponding industry-occupation cell. Figures are based on the dual connected sample, and include cells for which both women and men have zero fatalities.

## Table A1: Summary Statistics

| | Full Population | | FTPA Dominant Jobs | | Attached CS | | Dual CS | | Dual CS Stayers | | Dual CS Movers | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Women | Men | Women | Men | Women | Men | Women | Men | Women | Men | Women | Men |
| Age | 35.83 | 35.46 | 35.12 | 35.99 | 38.26 | 38.50 | 38.20 | 38.37 | 39.71 | 39.87 | 36.88 | 37.13 |
| Race White | 0.57 | 0.56 | 0.66 | 0.59 | 0.68 | 0.63 | 0.68 | 0.63 | 0.68 | 0.65 | 0.68 | 0.62 |
| Less than High School | 0.32 | 0.53 | 0.37 | 0.56 | 0.37 | 0.56 | 0.36 | 0.52 | 0.41 | 0.56 | 0.32 | 0.47 |
| High School | 0.41 | 0.33 | 0.44 | 0.35 | 0.39 | 0.33 | 0.39 | 0.36 | 0.37 | 0.33 | 0.41 | 0.38 |
| Some College | 0.05 | 0.03 | 0.06 | 0.03 | 0.06 | 0.03 | 0.06 | 0.04 | 0.06 | 0.03 | 0.07 | 0.04 |
| College or More | 0.21 | 0.10 | 0.13 | 0.06 | 0.18 | 0.07 | 0.18 | 0.09 | 0.16 | 0.07 | 0.20 | 0.10 |
| Contracted Weekly Hours | 39.46 | 42.03 | 42.54 | 43.30 | 41.77 | 43.00 | 41.63 | 42.79 | 41.43 | 42.60 | 41.79 | 42.95 |
| Log Hourly Wage | 1.33 | 1.42 | 1.22 | 1.41 | 1.52 | 1.64 | 1.54 | 1.72 | 1.54 | 1.77 | 1.55 | 1.68 |
| Total Experience (Years) | 18.35 | 18.98 | 18.44 | 20.22 | 21.90 | 23.23 | 21.80 | 22.89 | 23.62 | 24.72 | 20.19 | 21.38 |
| Job Tenure (Months) | 62.36 | 51.69 | 50.83 | 49.75 | 91.03 | 81.36 | 90.91 | 84.66 | 116.48 | 113.43 | 68.40 | 60.80 |
| Pooled Fatality Rate (per 100,000) | 0.02 | 0.07 | 0.03 | 0.08 | 0.02 | 0.08 | 0.02 | 0.07 | 0.02 | 0.07 | 0.02 | 0.07 |
| Gender-Specific Fatality Rate | 0.01 | 0.08 | 0.01 | 0.09 | 0.01 | 0.08 | 0.01 | 0.08 | 0.01 | 0.07 | 0.01 | 0.08 |
| Zero Fatality Rate | 0.26 | 0.15 | 0.33 | 0.11 | 0.37 | 0.12 | 0.37 | 0.13 | 0.36 | 0.14 | 0.38 | 0.13 |
| N | 134361238 | 194907785 | 40262142 | 70619734 | 9010984 | 20137854 | 8193244 | 14567312 | 3836088 | 6604350 | 4357156 | 7962962 |

Notes: FTPA Dominant Jobs sample includes jobs with the highest annual earnings for each worker between ages 23-65, with 30 or more contracted hours per week, excluding government jobs, temporary jobs, and jobs in 2-digit industry by 3-digit occupation cells that have fewer than 10,000 full-time full-year equivalent workers in the three-year moving average window used to calculate fatality rates. Attached CS sample includes the subset of the FTPA Dominant Jobs sample containing workers with a dominant job in each year between 2005-2010, who are employed at an establishment in the largest connected set corresponding to their gender. Dual CS contains the subset of workers in the Attached CS sample who were employed at establishments in both the largest male and female connected sets. Dual CS Stayers includes workers in the dual connected set who remain employed in the same establishment in all years, and Dual CS Movers is the complement of this set within the Dual CS sample. Fatality Rate is measured in deaths per 1,000 full-time full-year equivalent workers.

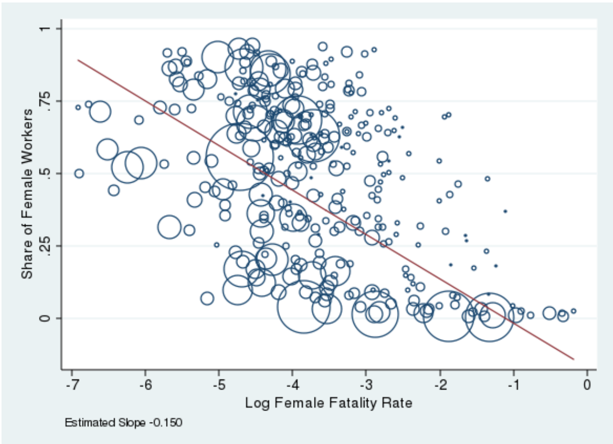
Figure A1: Distribution of Measurement Error in Female Fatality Rates



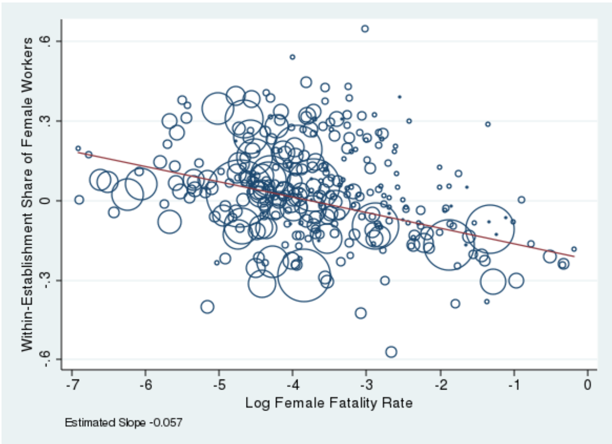

Notes: Figure plots the kernel density of the gender by 3-digit occupation by 2-digit industry fatality rate minus the gender-pooled 3-digit occupation by 2-digit industry fatality rate (solid line) and minus the female 2-digit average industry fatality rate (dashed line), all measured as three-year moving averages. The sample includes women in the attached dominant jobs sample with positive fatality rates, with fatality rates truncated at the 99th percentile for ease of presentation.

Figure A2: Female Worker Shares versus Log Fatality Rate

(a) Aggregate Distribution

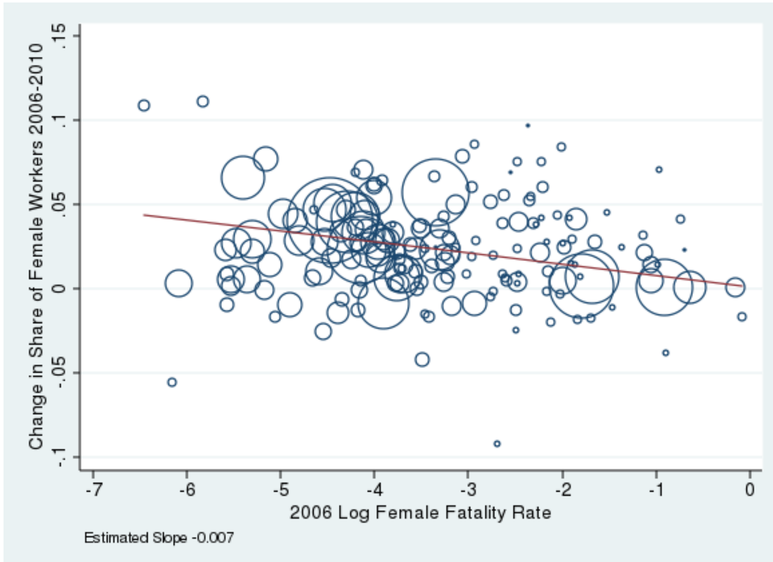



(b) Within-Establishment Distribution




Notes: Fatality rates are female-specific averages by 2-digit industry and 3-digit occupation pooled over all years from 2003–2010, measured in log of deaths per 1,000 FTFY workers. The sample includes the full population of women with positive fatality rates, with fatality rates truncated at the 99th percentile for ease of presentation. Circle sizes are proportional to the total number of workers in the industry-occupation cell.

Figure A3: Intertemporal Changes in Gender Shares by Fatality Rate



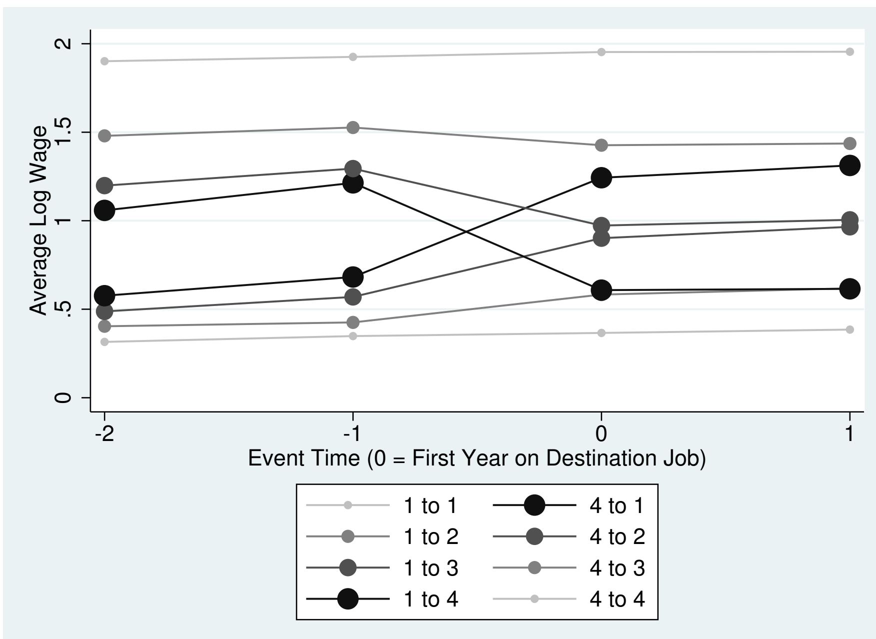

Notes: Vertical axis is the change between 2006-2010 in the employment-weighted share of female workers by 2-digit industry and 3-digit occupation. Horizontal axis is the log of the female-specific fatality rate in the 2-digit industry and 3-digit occupation in 2006, measured in deaths per 1,000 FTE worker-years. The sample includes the full population of workers with positive fatality rates, with fatality rates truncated at the 99th percentile for ease of presentation. Circle sizes are proportional to the total number of full-time workers of both genders in the industry-occupation cell.

Table A2: Components of Gender Wage Gap and Variance Decomposition

| | Women | Men | Gender Gap | |
|---|---|---|---|---|
| Log Wage | 1.543 | 1.726 | 0.183 | |
| First Stage Controls (Exp. and Year) | 0.668 | 0.650 | –0.018 | |
| Occupation Effects | 0.030 | 0.004 | –0.026 | |
| Person-Specific Component | –0.079 | 0.046 | 0.125 | |
| Establishment Assignment | 0.216 | 0.273 | 0.057 | |
| Variance Components | Women | | Men | |
| | Component | Variance Share | Component | Variance Share |
| SD of Log Wages | 0.74 | 100% | 0.69 | 100% |
| SD Worker Effects | 0.59 | 63% | 0.54 | 60% |
| SD Estab-Gender Effects | 0.44 | 35% | 0.36 | 27% |
| SD of $X\beta$ | 0.22 | 9% | 0.19 | 7% |
| SD Residual | 0.13 | 3% | 0.14 | 4% |
| Cov $(\theta, \Psi^g)$ | 0.04 | 7% | 0.04 | 9% |
| Cov $(\theta, X\beta)$ | 0.00 | 1% | –0.00 | -1% |
| Cov $(\Psi^g, X\beta)$ | –0.07 | -12% | –0.03 | -7% |

Notes: Estimates are based on Equations 1 and 2, estimated on the dual connected set, with normalized gender-establishment effects. Occupation effects include combined effect of differences in the distribution of men and women across 1-digit occupation cells, and the conditional average gender gap within occupations. $X\beta$ in the bottom panel includes the combined contribution of both first and second stage observed controls and fatality rates.

Figure A4: Event Study For Female Workers Who Change Jobs




Notes: The figure shows the average wage in each year surrounding a job change, for female workers in the dual connected set who change jobs and held their origin job for at least two years. Observations are grouped into cells based on quartiles of the origin and destination wage. The jobs are classified by their quartile in the female establishment effect distribution. The figure indicates no evidence for the most likely violations of exogenous mobility: (1) systematic declines in residual log wages prior to workers departing, and (2) asymmetric average changes in log wages associated with movements up versus down the distribution of establishment wage effects $\Psi^g_{J(i,t)}$.

Table A3: Gender Differences in Compensating Wage Differentials by Age

| Fatality Rate | | | |
|---|---|---|---|
| | (1) Men | (2) Women | (3) Both |
| Fatality Rate*Age 20s | .121* | 0.017 | 0.116* |
| | (0.002) | (0.008) | (0.002) |
| Fatality Rate*Age 20s*Female | | | –0.096* |
| | | | (0.009) |
| Fatality Rate*Age 30s | 0.205* | 0.215* | 0.203* |
| | (0.002) | (0.006) | (0.002) |
| Fatality Rate*Age 30s*Female | | | 0.014 |
| | | | (0.007) |
| Fatality Rate*Age 40s | 0.173* | 0.258* | 0.176* |
| | (0.002) | (0.007) | (0.002) |
| Fatality Rate*Age 40s*Female | | | 0.072* |
| | | | (0.008) |
| Fatality Rate*Age 50s | 0.093* | 0.131* | 0.091* |
| | (0.003) | (0.011) | (0.003) |
| Fatality Rate*Age 50s*Female | | | 0.050* |
| | | | (0.012) |
| Fatality Rate*Age 60s | –0.066* | –0.324* | –0.075* |
| | (0.007) | (0.039) | (0.007) |
| Fatality Rate*Age 60s*Female | | | –0.191* |
| | | | (0.042) |
| N | 13,985,793 | 8,131,646 | 22,117,439 |
| R-Sq | 0.958 | 0.970 | 0.971 |

Notes: Analysis sample includes dominant jobs of attached workers in the dual connected set. Column 5 includes both men an women, with gender-specific establishment effects and gender-specific coefficients on all other control variables (year, experience, and occupation). 'Fatality Rate' is measured in deaths per 1,000 FTFY workers and is Winsorized at the 99th percentile. Log wages are Winsorized at the 1st and 99th percentiles. * Indicates significance at the 0.01 level.

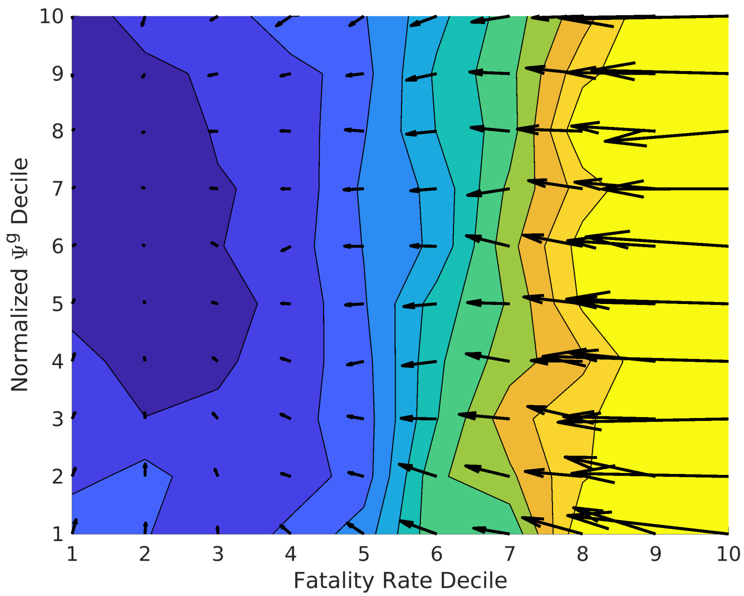
Table A4: Job Assignment and Gender Differences in Establishment Effects and Safety

| | All Jobs | Jobs Held by Women | Jobs Held by Men |
|---|---|---|---|
| Male Fatality Rate | 0.055 | 0.034 | 0.067 |
| Female Fatality Rate | 0.014 | 0.011 | 0.016 |
| Normalized Male Establishment Effect | 0.246 | 0.200 | 0.273 |
| Normalized Female Establishment Effect | 0.239 | 0.216 | 0.252 |
| Difference | 0.007 | -.016 | 0.021 |
| N | 18,632,474 | 6,896,311 | 11,736,163 |

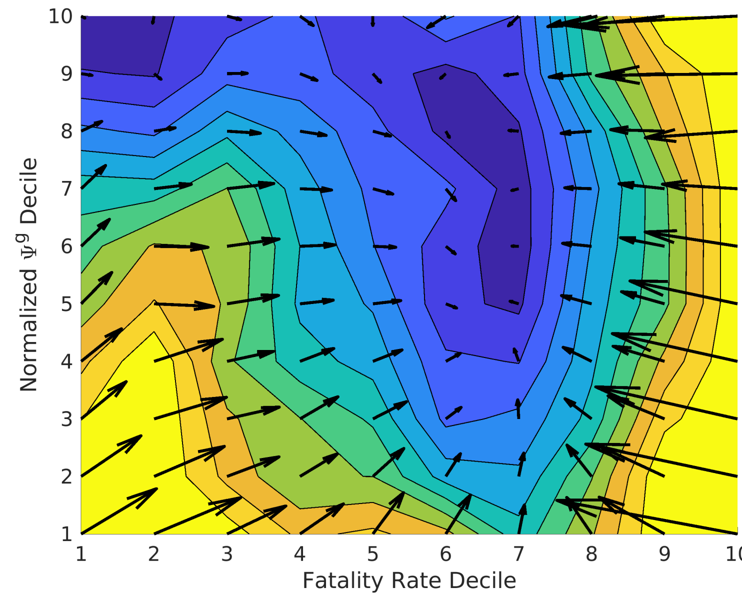
Notes: Estimates are based on the dual connected set. Column 1 reports averages for all jobs, as though all workers were either women or men. Column 2 reports averages among women relative their male counterparts at the same establishment (normalized establishment effects) or industry-occupation cell (fatality rates). Column 3 reports comparable averages for male workers. Gender-specific fatality rates are measured in deaths per 1,000 FTFY workers.

Figure A5: Job-to-Job Transition Gradient Fields
Establishment Pay versus Risk

(a) Women

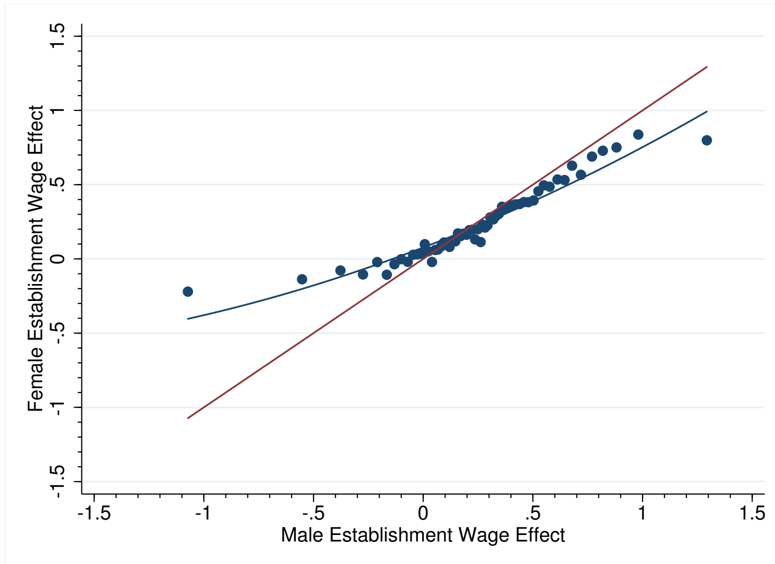



(b) Men



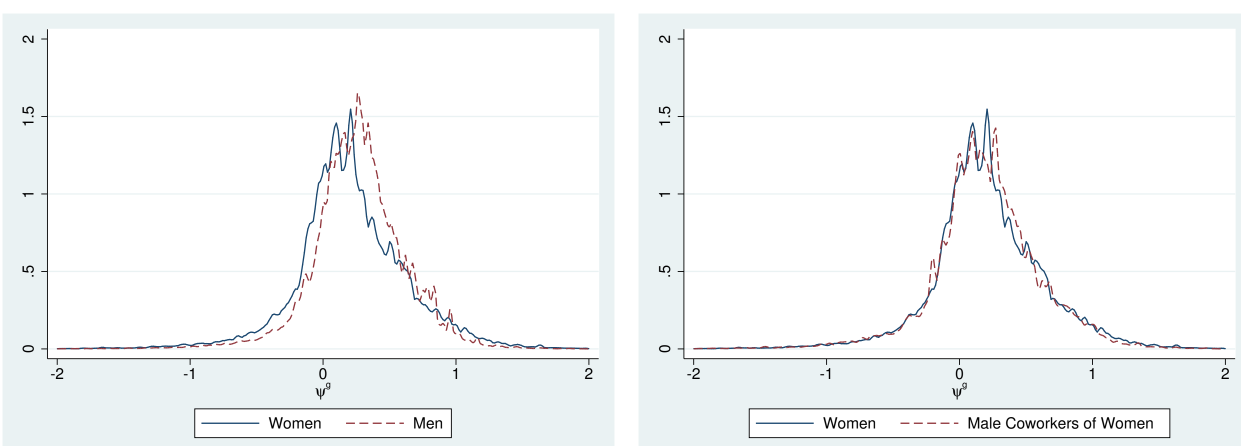

Notes: Samples include attached full-time women/men between ages 23-59 employed in the largest dual connected set of establishments, and excludes origin jobs with zero fatality rate. Deciles defined based on gender-specific wage and fatality rate distributions. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A6: Binned Scatterplot of Within-Establishment Male versus Female Normalized Establishment Wage Effects



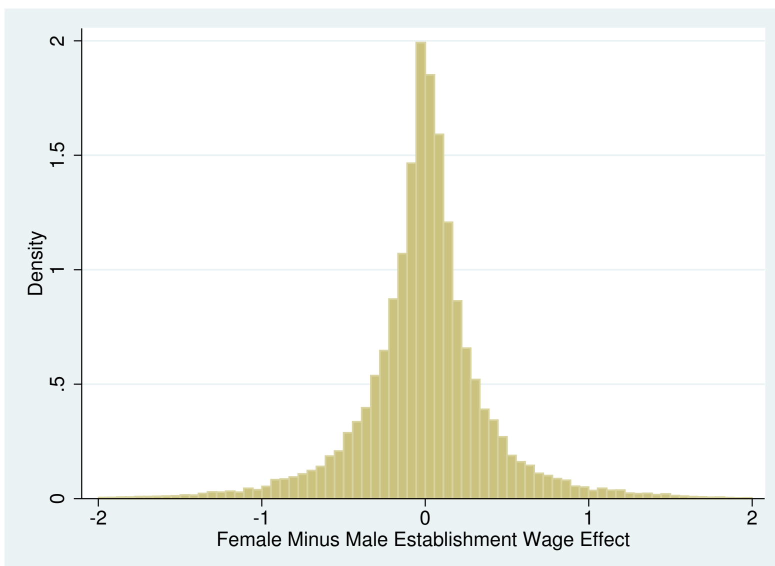

Notes: Figure depicts a binned scatterplot of the estimated female and male establishment effects for all jobs, normalized relative to the gender-specific average establishment effect in the lowest five industries. The red line is a 45-degree line, and the blue line is the predicted quadratic fit.

Figure A7: Establishment Wage Effects: Decomposing Assignment to Establishments and Treatment Conditional on Establishment



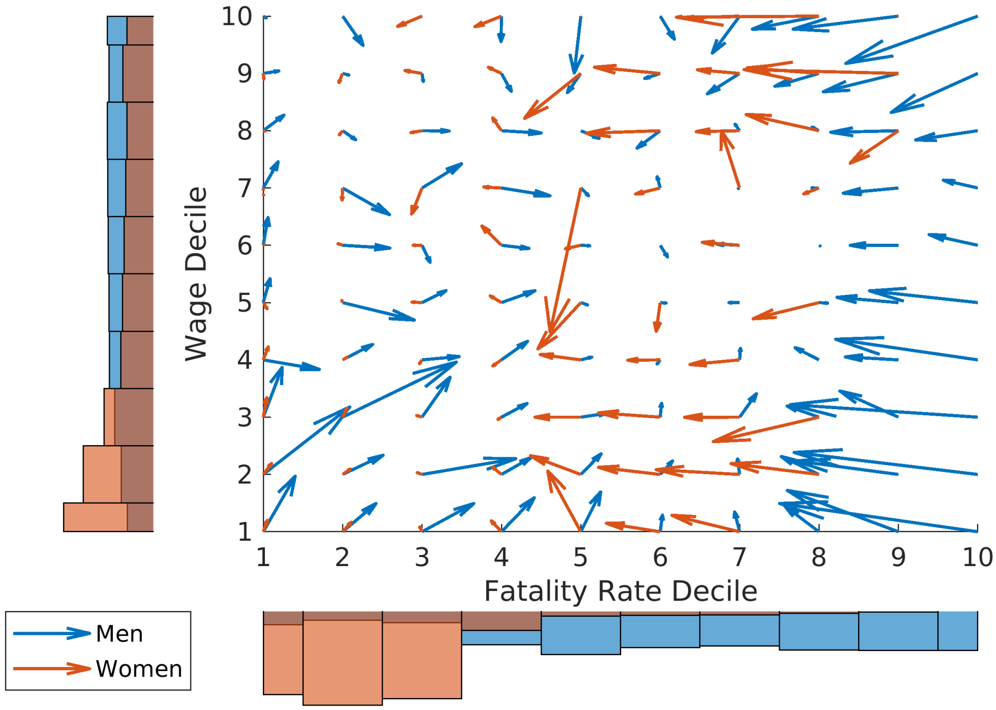

Notes: Subfigure A7a plots the kernel density of the normalized distribution of $\Psi^g_{J(i,t)}$ for women and men. Subfigure A7b plots the same kernel density for women, compared to the counterfactual kernel density that women would have received if they had been men employed at the same establishment. This counterfactual distribution is estimated based on the male coworkers of women, weighted by the distribution of jobs held by women.

Figure A8: Histogram of Difference between Estimated Female and Male Establishment Effects



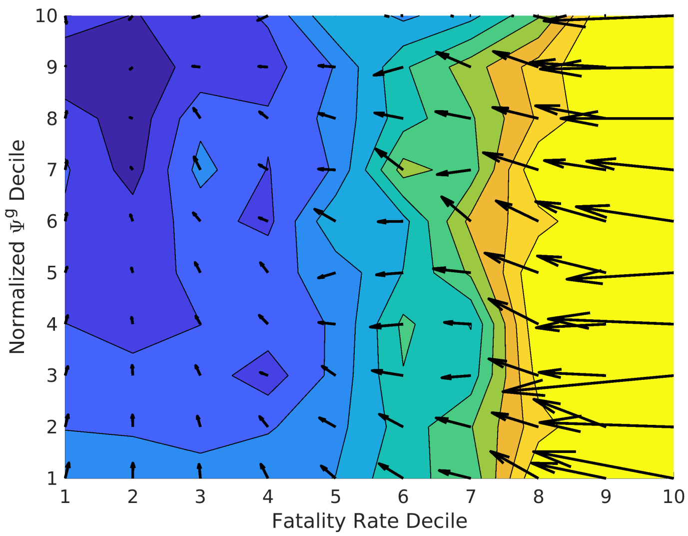

Notes: Distribution is based on normalized gender-specific establishment effects.

Figure A9: Job-to-Job Transition Gradient Field, Ages 60-65



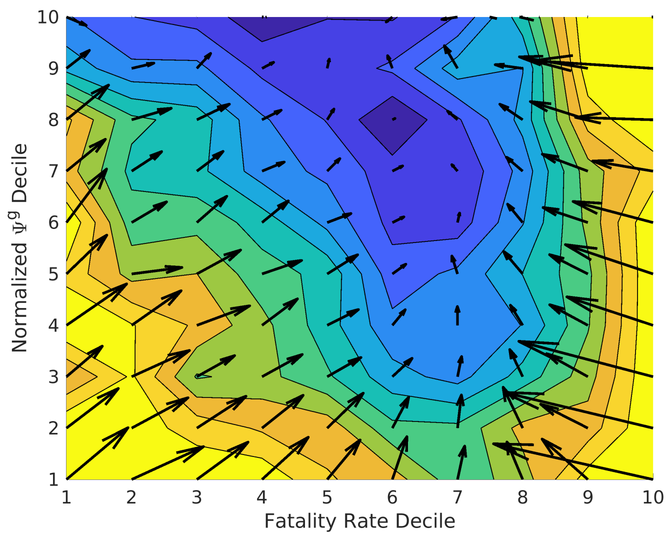

Notes: Sample includes attached full-time women and men between ages 60-65 employed in the largest dual connected set of establishments, and excludes origin jobs with zero fatality rate. Deciles defined based on combined female and male distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Marginal density functions depicted at origin jobs.

Figure A10: Female Job-to-Job Transition Gradient Field
Restricted to Separations Caused by Worker Resignation



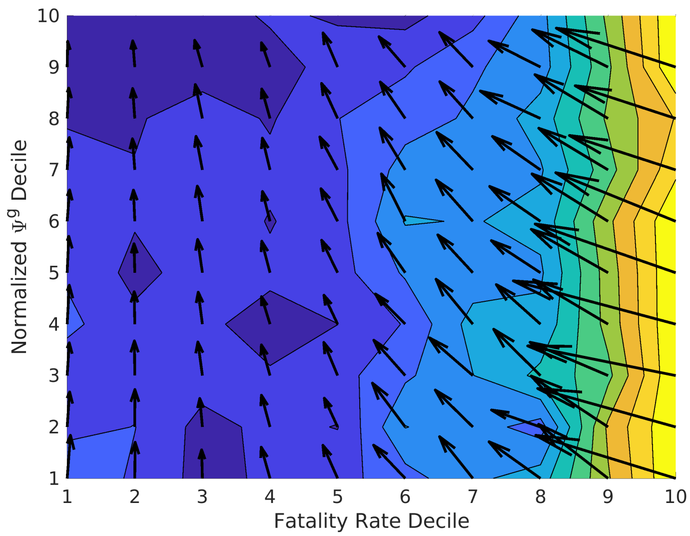

Notes: Sample includes attached full-time women between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes that originate with a worker resignation. Deciles defined based on female distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A11: Male Job-to-Job Transition Gradient Field
Restricted to Separations Caused by Worker Resignation



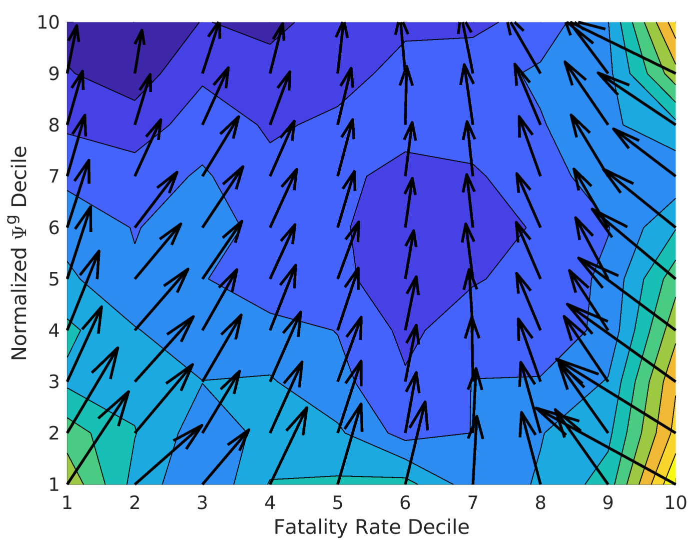

Notes: Sample includes attached full-time men between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes that originate with a worker resignation. Deciles defined based on male distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A12: Female Job-to-Job Transition Gradient Field
Conditional on Moving Up $\Psi^g$ Distribution



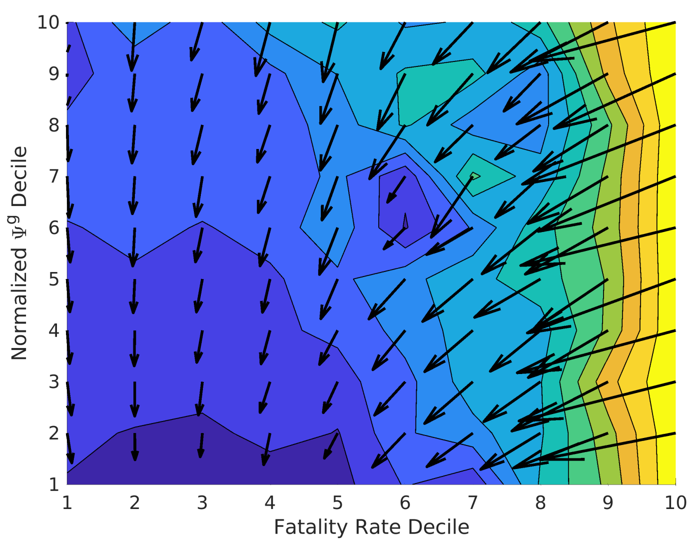

Notes: Sample includes attached full-time women between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes associated with an increase in the normalized $\Psi^g$. Deciles defined based on female distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A13: Male Job-to-Job Transition Gradient Field
Conditional on Moving Up $\Psi^g$ Distribution



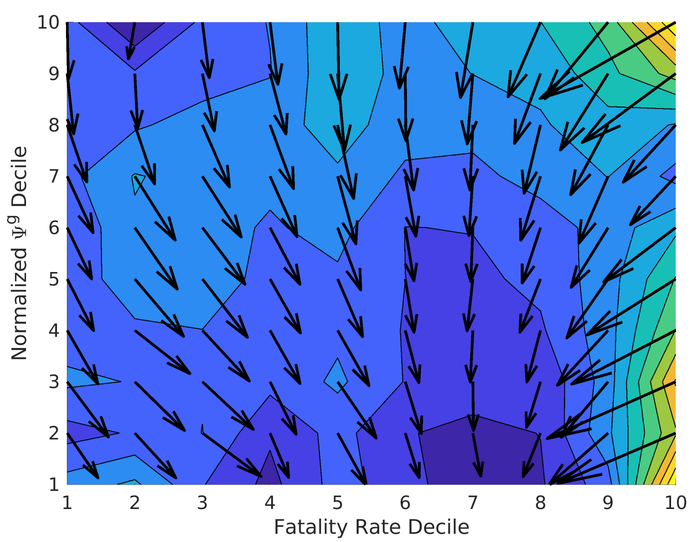

Notes: Sample includes attached full-time men between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes associated with an increase in the normalized $\Psi^g$. Deciles defined based on male distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A14: Female Job-to-Job Transition Gradient Field
Conditional on Moving Down $\Psi^g$ Distribution




Notes: Sample includes attached full-time women between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes associated with a decrease in the normalized $\Psi^g$. Deciles defined based on female distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.

Figure A15: Male Job-to-Job Transition Gradient Field
Conditional on Moving Down $\Psi^g$ Distribution




Notes: Sample includes attached full-time men between ages 23-59 employed in the largest dual connected set of establishments, excluding origin jobs with zero fatality rate, and including only job changes associated with an decrease in the normalized $\Psi^g$. Deciles defined based on male distributions, using gender-specific fatality rates. Vectors indicate directions and relative, not absolute, magnitudes of changes in wages and fatality rates associated with job-to-job transitions originating from each cell. Contours indicate level sets based on the relative lengths of vectors.