Policy Research Working Paper 10899

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Katharina Fietz
Aiko Schmeißer

WORLD BANK GROUP
Social Protection and Jobs Global Practice
September 2024

Policy Research Working Paper 10899

Abstract

This paper studies the impact of working with same-race coworkers on individuals' retention at firms. Using administrative employer-employee data from Brazil, the paper exploits unexpected deaths of workers from different racial groups as exogenous shocks to peer group composition. The findings show that a decrease in the non-white share of coworkers reduces the retention of non-white workers but does not affect the retention of white workers. The effects are driven by non-whites quitting and moving to new jobs with more peers of the same race than in their old jobs. The findings highlight how peer dynamics can contribute to racial segregation across workplaces.

This paper is a product of the Social Protection and Jobs Global Practice. It is part of a larger effort by the World Bank to provide open access to its research and make a contribution to development policy discussions around the world. Policy Research Working Papers are also posted on the Web at http://www.worldbank.org/prwp. The authors may be contacted at kfietz@worldbank.org and aiko.fiete.schmeisser@uni-potsdam.de.

Racial Peer Effects at Work: Evidence from Worker Deaths in Brazil *

Katharina Fietz † **Aiko Schmeißer** ‡

JEL Classification: J15, J21, J63

*We thank Marco Caliendo, Javier Feinmann, Katrin Huber, Simon Jäger, Jann Lay, Elizabeth Linos, Shushanik Margaryan, Suanna Oh, Amanda Pallais, Josefina Posadas, Nina Roussille, Pedro C. Sant'Anna, Sulin Sardoschau, Liam Wren-Lewis, and participants of the Development Economics Network Berlin (DENeB) Workshop, Ruhr Graduate School in Economics (RGS) Doctoral Conference, Berlin Network of Labor Market Research (BeNA) 20 Year Jubilee Conference, Doctorissimes PhD Conference, European Development Research Network (EUDN) PhD Workshop, Research Institute for Development, Growth and Economics (RIDGE) Workshop on Public Economics, German Development Economics (GDE) Conference, Warwick Economics PhD Conference, European Society for Population Economics (ESPE) Annual Conference, International Conference on Development Economics (ICDE), and the 39th meeting of the European Economic Association for valuable comments. In addition, we are grateful to the Brazilian Ministry of Labor for granting data access to the *Relação Anual de Informações Sociais* (RAIS).

†University of Göttingen, GIGA; katharina.fietz@giga-hamburg.de

‡University of Potsdam, BSoE; schmeisser@uni-potsdam.de

1 Introduction

Labor markets exhibit high levels of ethnic and racial segregation. Evidence from multiple countries consistently shows that minority workers are segregated from majority workers (Åslund and Skans, 2010; Forth *et al.*, 2023; Glitz, 2014; Hellerstein and Neumark, 2008). In Brazil, the focus of our study, about 80% of white workers' colleagues are also white, while for non-white workers the share of white peers is substantially lower, at about 50% (Gerard *et al.*, 2021). Leading explanations for labor market segregation include differences in residence, education, and job choices across racial groups (Hellerstein and Neumark, 2008), as well as differential hiring policies due to discriminatory behavior of managers or co-racial referral hiring (Giuliano *et al.*, 2009, 2011; Hsu Rocha and Dias, 2021; Miller and Schmutte, 2023). The role of post-hiring dynamics in explaining racial segregation has received less scrutiny, although the composition of a firm's workforce depends not only on who enters the firm but also on who stays.

In this paper, we examine one critical factor that can influence the retention of workers at a firm: the presence of same-race peers. A long tradition of theories in the social sciences, based on similarity attraction, social identity, social categorization, and discrimination (Byrne, 1971; McPerson *et al.*, 2001; Tajfel, 1981; Turner *et al.*, 1987), suggests that individuals exhibit homophilic preferences: People "love those who are like themselves", as already noted by Aristotle (1835) in his *Nichomachean Ethics*. If homophilic preferences based on race prevail in the workplace, a reduction in exposure to same-race peers may lead workers to move to a different job. The same prediction arises in related theories of organizational demography (Pfeffer, 1983). Reduced similarity along demographic characteristics can entail lower levels of cooperation and communication, hinder social integration within a group, and weaken mentoring relationships and informal support networks among coworkers (Hoffman, 1985; Ragins, 1997; Tsui and O'Reilly III, 1989). Thereby, a decrease in the presence of same-race peers may lower job satisfaction and, ultimately, reduce retention at the firm (Jackson *et al.*, 1991; O'Reilly III *et al.*, 1989; Riordan and Shore, 1997).

We study the effects of a sudden change in the racial makeup of coworker peer groups on worker retention. Using matched employer-employee data from Brazil, we exploit exogenous shocks to the composition of coworkers that arise from unexpected deaths of workers from different racial groups. We compare the retention of incumbent workers in peer groups that encounter the death of a non-white worker with those confronted with the death of a white worker. Causal identification thus relies on the race of the deceased worker being independent of all other factors that drive incumbent workers' retention. We check that, after we control for the number of non-white and white workers before the death, incumbent characteristics do not predict whether a deceased worker is non-white or white. Thus, when comparing peer groups with the same initial racial composition, the race of the deceased worker provides us with an exogenous shock to the racial composition among the remaining incumbent workers.[1]

Our analysis concentrates on small peer groups of no more than 30 workers in the same occupation within an establishment. We verify that the death of a non-white vs. white worker leads to a quantitatively meaningful and lasting shift in the racial composition of these peer groups. For our estimation strategy, we rely on simple cross-sectional regressions that relate incumbents' retention after the death to a dummy for whether the deceased worker is non-white, controlling for the pre-death racial composition and other incumbent characteristics. Using comprehensive administrative data from Brazil, this set-up allows us to study the causal effects of the racial coworker composition on workers' retention for a large set of firms.[2]

We find that experiencing the loss of a non-white coworker – relative to the loss of a white coworker – significantly reduces the retention of non-white incumbents. In quantitative terms, our results show that the likelihood of non-white incumbents staying at the same firm decreases by 1.1 percentage points (1.8% relative to the control mean) in the three years after the death. In contrast, we find no significant effects on the retention rates of white incumbents. Our results highlight that peer effects are not symmetric across racial groups: Non-white workers are more strongly affected by the racial composition of their peers than white workers. Several robustness checks, in which we vary the inclusion of control variables in the specification, estimate duration models, or use alternative peer group definitions, confirm our main results.

To shed light on the potential mechanisms driving the lower retention rates of non-white workers following the reduction in the share of same-race coworkers, we conduct several additional analyses. First, we show that the effects are driven by worker-initiated quits rather than employer-initiated layoffs. Given that in Brazil, workers who quit the job on their own are not entitled to severance payments and unemployment benefits, quits likely capture workers' labor supply driven by preferences for different job characteristics. Moreover, we do not find any significant effects on the earnings of non-white incumbents. The absence of effects on layoff probabilities and earnings suggests that employers' demand for incumbent workers is not affected by the loss of a same-race peer. Second, we track workers after they separate from their initial firm. For the subset of workers who start a new formal job in the same or the following

[1] Exploiting departures of workers due to deaths has another advantage: It rules out that the effects on the retention of incumbents are driven by racially segregated referral hiring, whereby departing workers may recruit their former coworkers to their new jobs (Hensvik and Skans, 2016; Miller and Schmutte, 2023).

[2] Jäger and Heining (2022) and Bertheau *et al.* (2022) leverage worker deaths as shocks to firms' overall labor supply to study replacement costs from worker exits. In contrast, we exploit the characteristics of the deceased worker which impact the composition of workers at a firm. Illing *et al.* (2023) examine how exogenous vacancies after worker deaths can lead to different labor market outcomes for male and female replacement workers. We, instead, analyze how deceased workers of different races affect the retention of incumbent workers.

year after their separation, we compare the characteristics of the old and the new job. We find that the death-induced decline in the non-white share of the old peer group leads non-white workers to move to new jobs that feature again a higher non-white share than the old jobs. Overall, these results are consistent with the idea that homophilic preferences for working with more same-race peers drive the initial separations of non-white workers.

Heterogeneity analyses reveal that the retention effects vary widely across different job types and peer group characteristics. We find larger effects in white-collar occupations than in blue-collar occupations, supporting the notion that the demand for workplace amenities – in our case, the preference to work with same-race peers – increases with income (Hamermesh, 1999; Pierce, 2001). Moreover, the effects are less pronounced in occupations requiring more teamwork, consistent with the contact hypothesis that predicts lower racial biases when group members frequently interact (Allport, 1954). In line with this, we also find that the impacts are smaller among incumbent workers who have a longer tenure in the firm alongside the deceased worker and thus had more opportunities to form a closer relationship with the deceased worker. In addition, the negative retention effects for non-white incumbents appear to be smaller in peer groups with a lower initial share of non-white workers. In these peer groups, non-whites have more contact with white coworkers which may reduce their in-group bias in coworker preferences.[3]

We also compare the effects of having same-race or same-gender coworkers. Changes in the gender peer group composition, resulting from the death of a female – as opposed to a male – coworker, are not found to affect the retention of female and male incumbents. However, we find important interactions: non-white women, i.e. the most disadvantaged group in the labor market, are most strongly affected by their peer group composition. Specifically, the retention of non-white females decreases when losing a same-race *or* same-gender coworker but is unaffected when losing a coworker that has the same race *and* the same gender. These results highlight that unbundling asymmetries across diverse race and gender categories is important to draw a comprehensive picture of coworker interactions at work.

Our analysis speaks to several strands of literature. First, we contribute to the literature on peer effects at the workplace. While economic studies have focused on the impact of peer skills and productivity (Cornelissen *et al.*, 2017; Herbst and Mas, 2015; Mas and Moretti, 2009; Messina *et al.*, 2023), interdisciplinary research has also examined the impact of the demographic composition of peers (for surveys, see Joshi *et al.* (2011) and Williams and O'Reilly III (1998)). The latter studies mainly use data from single (or a small number of) firms in high-income

[3] This result is also consistent with models of labor market sorting based on amenity preferences (Rosen, 1986): Non-white workers who have the weakest preferences for working with similar peers may sort into jobs with fewer non-white coworkers and also react less strongly to an unexpected decline in the proportion of same-race coworkers.

countries and often find that higher demographic similarity to coworkers is positively correlated with worker retention (Elvira and Cohen, 2001; Hirsch *et al.*, 2020; Leonard and Levine, 2006; Sørensen, 2004; Zatzick *et al.*, 2003). Causal evidence is provided by Linos *et al.* (2024), who exploit random team assignments within a high-wage professional services firm in the U.S. and find very similar results to ours: A larger share of white peers only increases the turnover of black women with no effects on other race and gender groups. We add to this literature by studying the effects of racial coworker composition on workers' retention using high-quality administrative data from Brazil. Exploiting exogenous changes in the peer group composition from unexpected worker deaths, we can isolate these changes from other confounding factors, such as firm management practices, and study the causal effects of racial coworker similarity for a large and diverse set of firms in a middle-income country.

Second, we speak to a large body of literature that studies cross-race interactions and racial segregation in various social contexts. Many papers have tested the "white flight" hypothesis to understand residential segregation, examining how neighborhood choices of whites are shaped by changes in the local exposure to minorities (Bayer *et al.*, 2024; Boustan, 2010; Card *et al.*, 2008; Shertzer and Walsh, 2019), and to understand school segregation, studying how white parents react to changes in the racial composition of schools (Baum-Snow and Lutz, 2011; Caetano and Maheshri, 2017; Cascio and Lewis, 2012; Reber, 2005). We complement this body of work by shedding new light on the retention effects of cross-race interactions at the workplace. In doing so, we also reveal important asymmetries in inter-group dynamics. While many studies focus on how the majority group reacts to increased contact with the minority group, our results indicate that the retention effects of working with same-race coworkers are not symmetric across racial groups and are more pronounced among non-white workers.[4] Moreover, we add an intersectional perspective to this literature by uncovering important differences in coworker composition effects across race and gender groups.

Lastly, our paper contributes to research on racial preferences and discrimination in the labor market. Much of the work focuses on hiring discrimination by employers, typically identified in correspondence studies (e.g., Bertrand and Mullainathan, 2004; Kline and Walters, 2021; Kline *et al.*, 2022) or by comparing hiring decisions of black and white managers (Åslund *et al.*, 2014; Benson *et al.*, 2023; Giuliano *et al.*, 2009, 2011; Hsu Rocha and Dias, 2021). The classical Becker (1971) model of discrimination predicts that market forces will counteract taste-based

[4]Our results are consistent with evidence on homophily in friendship networks among school children and adults. The tendency of individuals to form same-type ties at higher rates than predicted by random assortment – called "inbreeding" in the homophily literature – is often found to be more pronounced among racial minority groups. This tendency works against the opportunity structure of their potential network ties which is dominated by the majority group (Laumann, 1973; Marsden, 1988; McPerson *et al.*, 2001; Shrum *et al.*, 1988).

discrimination of employers. Because prejudiced employers give up profits by paying more for preferred-race workers, in a competitive model, they will be driven out of business and will therefore not give rise to racial workplace segregation in the long run. In contrast, discrimination among employees may help explain market segregation also in a competitive equilibrium. By analyzing how separations, in particular worker quits, respond to a change in the racial coworker composition, we identify workers' revealed preferences for working with more same-race coworkers. Together, our results highlight how coworker preferences can generate post-hiring peer dynamics that amplify racial segregation across workplaces.[5]

The rest of the paper is structured as follows. Section 2 gives an overview of the institutional background and the used data. Section 3 explains the empirical strategy, followed by Section 4 which presents our results. Section 5 concludes.

2 Background and Data

2.1 Social and Legal Setting

Racial inequalities. Brazil is marked by substantial racial division. Having received the highest number of forced African immigrants among countries in the Americas, the country was founded as a society based on race-driven slavery. Although Brazil did not establish explicitly racist institutions akin to the Jim Crow era in the U.S., racial differences remained highly salient after the end of slavery (Skidmore, 1992). Legal interventions to guarantee equal labor market opportunities emerged relatively recently, starting with the passing of the Constitution in 1988 and subsequent laws in 1989 and 1995 that prohibited racial discrimination in employment and wage determination. Despite legislative advances, significant disparities in labor market outcomes persist. Non-whites earn, on average, 30%-35% less than white workers. While a substantial portion of the wage gap can be attributed to observable factors such as education and occupation (Firpo *et al.*, 2021; Ñopo, 2012), differential hiring and pay-setting policies of firms also contribute to racial wage disparities (Gerard *et al.*, 2021). In addition, non-white workers in Brazil are more likely to be unemployed and underrepresented in leadership positions (IBGE, 2022).

Given the historical absence of legal segregation policies, Brazil's society developed a more fluid notion of race compared to the U.S. context. Racial categories are mainly characterized

[5] Our work is closely related to Hedegaard and Tyran (2018) who elicit coworker preferences in a field experiment with secondary school students in Denmark. Students are hired to work in pairs on a real, albeit short and simple, task and are asked to choose between coworkers of known productivity who have either a Danish-sounding or a Muslim-sounding name. The authors find that individuals are willing to give up about 8% of their earnings to avoid working with a coworker of a different ethnicity.

based on skin tone rather than as fixed traits determined by inheritance. Consequently, there exists more ambiguity in classifying race. Official statistics, and our main data source, divide individuals into five main racial categories: *branco* (white), *preto* (black), *pardo* (brown), *amarelo* (yellow), and *indigena* (indigenous). The most salient disparities in labor market outcomes are observed between *branco* individuals and the combined group of *preto* and *pardo* individuals, which together comprise 99% of the population. *Preto* and *pardo* workers are shown to exhibit relatively similar levels of education, employment, and wages (Garcia de Oliveira *et al.*, 1981; Gerard *et al.*, 2021; Miller and Schmutte, 2023; Silva, 2000). Thus, we pool *preto* and *pardo* into one "non-white" group and compare outcomes across white and non-white individuals, following a large strand of research on racial disparities in Brazil (e.g., Cornwell *et al.*, 2017; Derenoncourt *et al.*, 2021; Gerard *et al.*, 2021; Hirata and Soares, 2020; Miller and Schmutte, 2023; Telles, 2014).

Separation regulations. The Brazilian labor market exhibits a very high degree of labor turnover. Around 45% and 80% of terminated formal job spells have lasted less than one year and three years, respectively (Britto *et al.*, 2022). In our analysis, we distinguish between employer-initiated separations ("layoffs") and employee-initiated separations ("quits"). Layoffs account for roughly three-quarters of all separations, while quits cover most of the remaining part. If workers are laid off without a just cause, they are entitled to various benefits, such as severance payments and unemployment insurance (UI). Severance payments are paid by the employer, who contributes every month 8% of workers' earnings to a mandatory savings account (FGTS). Upon layoff, the worker can access the account's balance, as well as an additional penalty for the firm equal to 40% of the balance. UI benefits amount to 1 to 1.76 minimum salaries and are paid for 3 to 5 months.[6] In contrast, if workers quit the job themselves, they are not entitled to any benefits. Quits are thus more likely to be driven by the workplace preferences of employees.

Informality. Brazil has a relatively large informal sector that accounts for roughly 45% of all workers in 2015 (Derenoncourt *et al.*, 2021). Our analysis will focus on separations from formal jobs and measure the racial composition among all formally employed coworkers which we can observe in administrative employment data. Survey evidence shows that formality rates are quite similar across racial groups (Gerard *et al.*, 2021), which suggests that conditioning on formality does not pose major limitations for understanding racial disparities in labor market

[6] Van Doornik *et al.* (2023) provide evidence that formal layoffs can be used to extract rents from the UI system by showing that firms are more likely to lay off workers when they become eligible for UI benefits, but continue to employ them informally while on benefits and recall them when the benefit period is over.

outcomes.

2.2 Main Data Source: RAIS

Our analysis uses the *Relação Anual de Informações Sociais* (RAIS) that provides matched employer-employee data on the near universe of formal jobs in Brazil between 2006 and 2021. Employers must report annual information on all formal jobs in the last year to the Brazilian Labor Ministry (*Ministerio do Trabalho*, MTE) which uses the data for administering various social security programs. Firms' compliance with reporting requirements is high as they have to fear large penalties when failing to submit complete records. Each record in the data entails information on a contract between an establishment and a worker in a given year, including average monthly earnings, contracted hours, type of contract, occupation, the establishment's location and industry, and hiring and separation dates. We also observe the type of separation: layoff, quit, termination of temporary contract, death, retirement, or transfer. Individual and establishment identifiers allow us to track separations of workers from an establishment.[7]

Moreover, the data includes worker characteristics, i.e. their age, gender, education, and race, as reported by the employer. Cornwell *et al.* (2017) document that a substantial share of workers have different races reported by different employers and that changes in race are not entirely random as they are correlated with wage changes upon job transition. To address these inconsistencies, we follow Gerard *et al.* (2021) and assign each worker the modal race as reported in all their contracts within the RAIS across all years.

3 Empirical Strategy

The racial composition of the workforce is likely influenced by a variety of firm and job characteristics, such as firms' management practices, hiring policies, training and promotion opportunities, and local labor market conditions, which could also affect worker retention. Omitting such unobservable characteristics would thus bias the retention effects of coworker similarity. Focusing on changes in the workplace composition over time by controlling for workplace fixed effects in longitudinal data (Hirsch *et al.*, 2020; Leonard and Levine, 2006), may not be enough to mitigate these concerns, as these changes may also be driven by changes in confounding firm and job characteristics.

We exploit exogenous changes in the racial coworker composition that arise from unexpected deaths of workers. Specifically, we compare the responses of incumbent workers in peer groups that experience the death of a non-white worker with those that experience the death of a white

[7] Throughout the paper, we use the term 'firm' interchangeably to refer to establishments.

worker. In the following, we describe the identification of deceased workers and incumbent workers in the administrative data. We also provide evidence that, when comparing peer groups with the same pre-death racial composition of employees, peer group characteristics do not predict whether a white or non-white worker dies. In addition, we verify that the death of a non-white worker, compared to the death of a white worker, leads to a lasting drop in the non-white share of the peer group. For these reasons, we can treat the race of the deceased worker as a conditionally exogenous and significant shock to the peer group's racial composition and can estimate its effect on incumbent workers' retention at the firm.

3.1 Sample Construction

Unexpected worker deaths. We identify worker deaths based on the employer notifications in the RAIS data. When employers report contracts that were terminated because the employee died, they distinguish between deaths due to a work accident, deaths while commuting to work, and all other deaths. We exclude work accident and commuting deaths in order to avoid concerns that safety conditions at and on the way to work may affect both the likelihood of worker deaths and the retention decisions of workers. Moreover, to rule out spurious death notifications, we drop deaths of individuals for whom we observe an employment contract at least 30 days the date of death (less than 3% of all deaths).

In order to identify separations that are likely unexpected for coworkers, we focus on deaths of workers younger than 65 who have a full-time (with at least 30 hours) and permanent job. Moreover, we only include individuals with a minimum tenure of three months at the time of death, as a large share of contracts are terminated in the first three months when firms can dismiss workers without cost. We also exploit employer reports on workers' sick leave periods and drop deaths of individuals who had any sick leave in the two years before the death (37% of all deaths).[8] While we do not observe the exact cause of death, this restriction allows us to exclude deaths associated with prolonged illness, such as cancer, and instead focus on unexpected deaths, such as those caused by accidents, strokes, or homicide. We consider deaths between 2009 and 2018, given that sick leave information started to be reported in 2007 and that we seek to follow incumbents for three years after the death.

Incumbent workers. We define the deceased worker's peer group as all incumbent workers at the time of death who work full-time at the same establishment and in the same 4-digit occu-

[8] All formal employees in Brazil have a right to 15 days of sick leave during which the employer continues to pay wages. After that, employees can receive sick leave benefits from the *Instituto Nacional do Seguro Social* (INSS) if they have contributed to the INSS for a minimum of 12 months. As every employee is entitled to the sick leave period offered by their employer, we can track all instances of employees taking sick leave.

pation as the deceased worker. The Brazilian classification of jobs – the *Classificacao Brasileira de Ocupacoes 2002* (CBO) – contains 611 occupational titles at the 4-digit level. Examples are civil engineers, secondary school teachers, telecommunication technicians, tourist guides, heavy machinery assemblers, and aircraft maintenance mechanics. Defining the peer group at the 4-digit level ensures that workers in the peer group can interact, collaborate, and learn from each other, which can affect their productivity, social integration, job satisfaction, and ultimately, their retention decisions (Messina *et al.*, 2023). Moreover, workers in the same occupation perform similar tasks and should thus be affected to the same extent by a potential increase in labor demand of firms that cannot easily replace the deceased worker (Jäger and Heining, 2022). We will test the robustness of our results when defining peers at the 6-digit level, which covers a total of 2,458 occupations. For example, within the 4-digit category "civil engineer", a 6-digit code differentiates various types of civil engineers, such as geotechnics engineers, hydraulics engineers, bridges and viaducts engineers, railway and metroways engineers, road engineers, and environmental engineers. Being able to identify peers in such narrow occupations helps us to focus on teams where non-white and white workers work closely together and share similar tasks.

Sample Restrictions. We restrict the sample to small peer groups to ensure that the death of one worker has a quantitatively meaningful effect on the racial coworker composition. In our baseline specification, we focus on peer groups with a maximum of 30 workers before the death, and in robustness analyses, we study even smaller peer groups with at most 20 or 10 workers. We additionally exclude peer groups that have multiple unexpected worker deaths in a given year such that our estimates are always driven by the loss of one worker. We also restrict the sample to "surviving peer groups" in which we observe at least one worker in every year of our observation window, i.e. three years before to three years after the death, allowing us to study a balanced panel of peer groups. Finally, we focus the analysis on workers in private-sector firms.

Summary statistics. We end up with a sample of 48,676 deceased workers whose peer groups include 413,061 incumbent workers. Thus, on average, there are 8.5 incumbents per peer group.[9] Table 1 reports summary statistics for the characteristics of deceased and incumbent workers. Compared to incumbents, deceased workers are older and more likely to be male and have more tenure, higher earnings, and less education. 36% and 38% of the deceased and incumbent workers, respectively, are non-white. Non-white incumbents have on average 0.5 years less

[9]Our peer groups are somewhat smaller than in other studies on workplace peer effects: for example, the average peer group size is 9 in Cornelissen *et al.* (2017), 12 in Hong and Lattanzio (2022), 28 in Linos *et al.* (2024), and 30 in Messina *et al.* (2023). While these papers use variation in the composition of all peers in a peer group, we exploit the death of only one coworker which has a larger effect on the composition of smaller peer groups.

education, earn about 20% lower wages, and have about 8 months less tenure at the firm than white incumbents.

Our main outcome variable is incumbents' retention at the establishment. Only about 50% of the incumbents in our sample remain in the establishment three years after the death. Figure 1 reports estimated Kaplan-Meier curves, depicting how retention rates evolve over time and vary by race of the incumbent. In line with evidence from the U.S. (Linos *et al.*, 2024; Sorkin, 2023), we observe a racial gap in retention rates: non-white workers have a 4.6 percentage points lower survival probability than white employees at the end of our 3-year observation period. Appendix Figure 4 distinguishes between employer-initiated separations ("layoffs") and employee-initiated separations ("quits"). More than 40% of incumbents are laid off three years after death, while only about 10% of jobs are quit by incumbents. Interestingly, we find that the racial gap in overall turnover is entirely driven by layoffs which are more likely to occur among non-whites. In contrast, non-white workers quit less often than white workers.[10]

3.2 Identification and Estimation

Our empirical strategy exploits shocks to the racial composition of incumbents' coworkers that arise from the death of a non-white vs. white coworker. Causal identification relies on the race of the deceased worker being independent of all other factors that drive incumbent workers' retention.

Balance checks. To examine the exogeneity of the deceased worker's race, we start by regressing a dummy for the deceased worker being non-white on various observable characteristics of incumbent workers, such as race, gender, education, age, wages, tenure, and firm size. Results are reported in Table 2, separately for all incumbents, non-white incumbents, and white incumbents. As can be seen in columns (1), (3), and (5), several characteristics are significantly correlated with the race of the deceased worker. A major difference between peer groups in which a non-white or white worker dies lies in their initial racial composition: as expected, there is a one-to-one relationship between the initial non-white share in the peer group and the likelihood of the deceased worker being non-white (see Appendix Figure A.2). Due to racial workplace segregation, the race of the deceased worker is also highly correlated with the race of incumbent workers. Column (1) shows that non-white incumbents are 36 percentage points more likely to be in a peer group with a non-white deceased worker.

In columns (2), (4), and (6) of Table 2, we flexibly control for the initial racial composition

[10]Search models with employer discrimination predict lower quit rates among racial minorities due to their lower outside options and higher search costs (Black, 1995; Whatley and Sedo, 1998).

before the death by including dummies for the exact number of non-white and white workers in the peer group before the death.[11] Importantly, this entirely eliminates the imbalance in incumbent characteristics. When separately considering non-white and white incumbents (as we will always do in our analysis), all incumbent characteristics are no longer significant (at the 5% level) in predicting whether the deceased worker is non-white or white.[12]

Apart from the variables shown in Table 2, we have also run regressions that add dummies for the state, occupation (1 digit), and sector (1 digit) of the incumbent. These dummies do also not jointly predict the race of the deceased worker (p-values = .224 for all incumbents, .252 for non-white incumbents, and .405 for white incumbents). In addition, we have assessed the balance of local-level mortality rates. As shown in Appendix Table A.1, the municipality's population share of deaths and deaths due to homicides or firearms, as well as their racial composition, are also not related to the race of the deceased worker after controlling for the initial number of non-white and white peers in the peer group.[13]

Estimating equation. For our empirical strategy, we exploit the conditional independence of the deceased workers' race: we compare peer groups with the same initial racial composition that are exposed to the death of a non-white vs. white worker. For this, we estimate simple cross-sectional regressions of the following form:

$$\begin{aligned} Y_{ijt}^r = \alpha_t^r + \beta_t^r \mathbb{1}[\,\text{Deceased} = \text{non-white}]_j \\ + \sum_k \sum_l \gamma_{klt}^r \mathbb{1}[\#\text{NW} = k, \#\text{W} = l] + \delta_t^r X_{ij} + \epsilon_{ijt}^r, \end{aligned} \tag{1}$$

where Y_{ijt}^r denotes an outcome of incumbent i from peer group j in year after death $t = \{-3, -2, ..., 3\}$. The regressor of interest, $\mathbb{1}[\,\text{Deceased} = \text{non-white}]_j$, is a dummy for whether the deceased worker is non-white. In order to restrict comparisons to peer groups with the same initial racial composition, as in Table 2, we include interacted dummies that flexibly control for the number of non-white workers k and white workers l in the peer group before the death. X_{ij}

[11] We measure the number of workers before the death as the number of incumbent workers on the day of death plus the deceased worker (and we differentiate incumbents and deceased workers by race). Panel B of Appendix Figure A.2 depicts the variation in the race of the deceased worker that remains when considering peer groups with the same initial racial composition.

[12] The only variable in Table 2 that remains significant in explaining the deceased worker's race is the incumbent's race: in peer groups with the same number of non-white and white workers before the death, incumbent workers are less likely to be non-white if the deceased worker is non-white (column (2)). Note that this is the first-stage effect that we want to exploit. Comparing peer groups with the same initial racial composition, the death of a non-white vs. white worker leads to a drop in the non-white share among the remaining workers. In section 3.3, we verify that there is also a lasting effect on the racial composition among all workers in the peer group, i.e. including incumbents and new hires.

[13] In Table 2, we only condition on the racial composition among workers in the peer group, i.e., in the same 4-digit occupation of the deceased workers. We have also verified that the racial composition among all workers in the establishment is balanced after we control for the racial composition in the peer group.

is a vector of additional control variables. In our baseline specification, we include dummies for the year of death, all incumbent characteristics shown in Table 2, as well as state $\times$ occupation (1 digit) $\times$ industry (1 digit) fixed effects to flexibly control for labor market conditions. Standard errors are clustered at the peer group level.

We estimate model (1) separately for each event period $t = \{-3, -2, ..., 3\}$ and for incumbents of race $r = \{\text{non-white, white}\}$. The pre-death estimates serve as additional balance checks to verify that the deceased workers' race is not related to when incumbents joined the firm. The post-death periods yield the estimates of interest that indicate the dynamic effects on incumbents' retention by year after death. We also estimate the average effect across all post-death periods by running model (1) jointly on all $t = \{0, 1, 2, 3\}$.

3.3 Effects on Racial Composition of Peer Group

Before turning to the results for retention outcomes of incumbent workers, we seek to verify that we can indeed view the death of a non-white vs. white worker as a significant shock to the racial composition of the peer group. If the deceased worker was immediately replaced by a new worker of the same race, the death would have no lasting impact on the make-up of the peer group. To examine the dynamic effects of a non-white vs. white death on outcomes at the peer group level, we estimate the following event-study model:

$$\begin{aligned} Y_{jt} = \alpha_j + \sum_{t \neq -1} \beta_t \mathbb{1}[\,\text{Deceased} = \text{non-white, Period} = t]_{jt} \\ + \sum_k \sum_l \sum_t \gamma_{klt} \mathbb{1}[\#\text{NW} = k, \#\text{W} = l, \text{Period} = t] + X_{jt} + \epsilon_{jt}, \end{aligned} \tag{2}$$

where Y_{jt} denotes an outcome of peer group j in year after death t. The effects of interest are captured by β_t which are the coefficients for interactions between a dummy for whether the deceased worker is non-white and dummies for each event year. For $t \geq 0$, they identify the dynamic treatment effects (relative to the year before death $t = -1$), and for $t < -1$ they check whether the outcome evolved in parallel in the pre-treatment period. α_j denotes peer group fixed effects that capture all time-invariant differences between peer groups with a non-white vs. white death. Similar to model (1), we include interacted dummies that allow trends to differ flexibly by the number of non-white workers k and white workers l in the peer group before the death.[14]

[14] The vector X_{ij} also adds fixed effects for state $\times$ occupation (1 digit) $\times$ industry (1 digit) groups and for the year of death, all interacted with event time t. Importantly, with the latter fixed effects our identifying variation only comes from comparing changes across peer groups with non-white and white deaths that occurred in the same year. Thereby, it avoids "forbidden comparisons" between late and early-treated groups that may lead to negative weights when averaging potentially heterogeneous, cohort-specific treatment effects in staggered DiD

Figure 2 shows the results for the number of workers (total and by race) and the non-white share of all workers in the peer group. We find that the death of a non-white vs. white worker leads to a significant and lasting shift in the racial composition of the peer group. The non-white share drops by almost 20 percentage points in the year of the death and then recovers only slowly over the three years after the death, after which it is still about 11 percentage points lower than before the death. This effect is driven by almost symmetrical increases in the number of white workers and decreases in the number of non-white workers in the peer group. In contrast, the total number of workers is not affected differently by the death of a non-white or white worker. Thus, with our empirical design, we exploit a shock to the racial composition but not to the size of the peer group.

In Appendix Figure A.3, we also explore the role of hiring in driving these effects. Interestingly, we find that the death of a non-white vs. white worker reduces the non-white share of new hires in the peer group by about 2 percentage points. In other words, following the loss of a worker of a given race, employers become less likely to hire new workers of the same race. This result is not consistent with the potential concern that non-white and white workers may perform different tasks even within our narrowly defined peer groups, which should make it more likely that deceased workers are replaced by new hires of the same race. Instead, preferences of job seekers to work with same-race coworkers may explain why employers are less likely to hire non-white workers after a negative shock to the non-white share of the existing workforce.

4 Results

4.1 Main Results

We now turn to our main findings for the effect on incumbent workers. Figure 3 reports the results from model (1) for incumbents' retention, i.e., their likelihood of working in the same establishment as at the time of death. Looking first at the period leading up to the worker's death, we do not observe significant differences in when incumbent workers have joined the firm depending on the race of the deceased worker. These results serve as an additional balance check consistent with the exogeneity of the deceased worker's race.

After the death, the retention outcomes of non-white and white incumbents start to diverge. Non-white incumbents' likelihood of staying in the same establishment decreases in the year of the death and remains lower over the subsequent three years. On average across all post-death

settings such as ours (De Chaisemartin and d'Haultfoeuille, 2020; Goodman-Bacon, 2021; Sun and Abraham, 2021). This stacking approach, first implemented by Cengiz *et al.* (2019), is robust to heterogeneous treatment effects.

periods, we estimate that the retention probability is reduced by about 1.1 percentage points, which is a drop of 1.8% relative to the sample mean. In contrast, for white incumbent workers, we do not find significant retention effects in any post-event year. In summary, the death of a non-white worker – relative to the death of a white worker – is found to reduce the retention of non-white incumbents but does not affect the retention of white incumbents.[15] The effects of a change in the racial coworker composition do not appear symmetric across racial groups, which is consistent with recent evidence from Linos *et al.* (2024).

Quantification. To facilitate the quantitative interpretation of our main results and to compare them to previous studies, we scale our estimates by the reduction in the non-white coworker share that incumbents experience due to the death of a non-white vs. white worker. For that, we first re-estimate model (1) using as the dependent variable the (leave-out) non-white share of an incumbent's coworkers in the initial peer group. Compared to the results at the peer group level described in Section 3.3, which give equal weight to small and large peer groups, this incumbent-level estimation provides us with an estimate of the effect that is representative across incumbents. As shown in Appendix Figure A.4, losing one non-white vs. white worker reduces the non-white share of incumbents' coworkers by about 6.8 percentage points on average over the three years after the death. The effect size is the same for non-white and white incumbents. Thus, our results imply that a 10 percentage point increase in the non-white share would increase the retention of non-white workers by about 1.6 percentage points. The retention of white workers, in contrast, is not significantly affected by the non-white share.

Our estimated retention effects for non-whites are somewhat smaller in magnitude compared to most of the existing workplace demography literature.[16] We interpret this finding as plausible since we capture a wide range of workers throughout different sectors, occupations, and skill levels in Brazil, whereas the existing evidence mainly comes from a small set of firms, in particular large and high-paying firms in the United States. Studies of low-wage workers and more representative samples of firms tend to find smaller effect sizes.[17]

Labor supply vs. labor demand responses. In Figure 4, we show that the retention effects

[15] In Appendix Table A.3, we show that the effects on non-white and white incumbents are significantly different from each other.

[16] Linos *et al.* (2024) find that a 10 percentage point increase in the share of white coworkers leads to a 4.6 percentage point increase in the probability of black workers leaving the firm two years after hiring. Zatzick *et al.* (2003) find that a 10 percentage point increase in the proportion of same-race coworkers is related to a decrease in voluntary turnover rates by 7.4%. Sørensen (2004) finds that one additional coworker of the same race correlates with a 3.5% lower probability of new hires voluntarily separating from the firm.

[17] In a sample of part-time, unskilled workers in a U.S. service firm, Leonard and Levine (2006) find that a 10 percentage points higher share of same-race coworkers is related to a 0.8 percentage point lower turnover rate. Within a representative sample of West German manufacturing firms, Hirsch *et al.* (2020) find a 2.2% increase in the job-to-job turnover rate for a 10 percentage point higher share of same-nationality coworkers.

on non-white incumbents are driven by quits and not by layoffs. The death of a non-white vs. white worker increases non-white incumbents' likelihood of having quit by 0.6 percentage points and does not significantly affect the likelihood of being laid off. The effect on quits is substantial amounting to 10% relative to the low sample mean of 6.0%. The results indicate that the death of workers from different racial groups can change employees' labor supply at the establishment. Non-white workers voluntarily separate from the peer group subsequent to the departure of a similar coworker, which is consistent with employees' preferences about the racial composition of their coworkers explaining the observed retention effects.

In contrast, employers' demand for incumbent workers does not appear to be affected by the death of a non-white vs. white worker. Labor demand effects could occur for several reasons. First, the racial coworker composition may impact worker productivity. If cooperation, communication, and learning are more effective in more homogeneous teams (Hjort, 2014), a lower share of non-white coworkers may decrease the productivity of non-white workers and, thus, employers' demand for these workers. Second, employers may also have preferences about the composition of their workforce. If employers value racial diversity (Dobbin and Kalev, 2022), they may seek to counteract the decline in the non-white share by increased efforts to retain non-white incumbents. Third, if workers of different races carry out different tasks even within our narrowly defined peer groups, labor demand effects may also arise from differential substitutability. For example, when a non-white worker dies, employers who cannot easily replace the deceased worker may increase their demand for incumbent non-white workers who perform similar tasks (Jäger and Heining, 2022). Overall, a change in the demand for non-white incumbents should lead to a change in their layoff probabilities, but we find no evidence of this. Moreover, labor demand effects should be reflected in an increase in earnings for non-white incumbents. As shown in Appendix Figure A.5, we also find no effects on earnings of incumbent workers.

4.2 Robustness

Next, we conduct a series of robustness checks to assess the sensitivity of our main results from Figures 3, 4, and A.5.

Specification. Given that our identification strategy relies on the conditional independence of the deceased worker's race, in Appendix Table A.2 we check that our results are robust to using alternative sets of control variables. In all specifications, we control for the initial number of non-white and white workers in the peer group and for the year of death. We vary whether or not we include controls for incumbent characteristics, state $\times$ occupation (1 or 2 digit) $\times$

industry (1 or 2 digit) fixed effect, or additional controls for local mortality rates. Our main results remain similar when subsequently removing or adding these control variables, which again underlines the exogeneity of whether the deceased worker is non-white or white.

In addition, we account for potential unobserved heterogeneity across peer groups with a non-white or white deceased worker. For that, we estimate a version of model (1) that pools non-white and white incumbents and includes interaction terms between all explanatory variables in the model and a dummy for whether the incumbent is non-white. In this model, we can include peer group fixed effects and still identify the differential effect of a non-white vs. white worker death on non-white vs. white incumbents. Results, reported in Appendix Table A.3, show that the effects on retention are significantly lower for non-white incumbents than for white incumbents and that this interaction effect remains almost unchanged when adding peer group fixed effects to the model.

Duration models. Instead of OLS, we also estimate duration models that take into account censoring in survival data. Specifically, we run Cox proportional hazard models to estimate how separation hazards differ in peer groups where the deceased worker is non-white or white. We include the same set of control variables and fixed effects as in model (1) and also cluster standard errors at the peer group level. Results are reported in Appendix Table A.4 and mimic our OLS results. The death of a non-white vs. white worker increases the hazard of separating from the firm by 2.1% for non-white incumbents. The duration model also confirms that the effects are driven by quits rather than layoffs. Again, we find no significant effect on separations of white incumbents.

Peer group definition. In any study of peer effects, it is crucial to carefully evaluate the suitability of the chosen peer group delineation. In our case, we seek to capture workers in an establishment that perform similar tasks and actually interact with each other, such that the death of a worker constitutes a meaningful change to the composition of their peer group. Thus, in Appendix Table A.5 we check the sensitivity of our results when narrowing the peer group definition. First, we consider incumbents who work in the same 6-digit occupation as the deceased worker, instead of the same 4-digit occupation as in our baseline approach. With almost 2,500 different occupation titles, the 6-digit level captures detailed occupational profiles that precisely organize the variety of tasks performed. Still, the results align very closely with our main findings. Second, we modify the peer group size considering peer groups with a maximum of 20 or 10, instead of 30, workers before the death. For non-white workers the results remain very similar when focusing on smaller peer groups. Interestingly, in peer groups with a maximum of 10 workers we also find significantly positive effects on the retention of white incumbents. This suggests that in small peer groups, where relationships among coworkers may be deeper, losing a coworker of a different race may make also white workers more likely to stay at the firm.

4.3 Heterogeneity

We will now exploit our diverse sample of workers and peer groups to investigate potential heterogeneities in how incumbent workers' retention is affected by losing same-race coworkers.

Occupation characteristics. We start by examining whether effects differ across occupation characteristics. First, we split our sample into white-collar and blue-collar occupations.[18] The nature of coworker interactions may differ substantially between these occupations. Moreover, if the demand for workplace amenities – specifically, the preference for same-race coworkers – increases with income (Hamermesh, 1999; Pierce, 2001), we expect the effects to be more pronounced in white-collar occupations. Second, we differentiate workers by the importance of teamwork in their occupation, using O*NET data on occupational task descriptions.[19] Two contradictory mechanisms could be present here: On the one hand, one could expect that in jobs requiring more teamwork, coworkers interact more strongly and thus may be more affected by the racial composition of their colleagues. On the other hand, the contact hypothesis would predict that racial biases are mitigated when workers interact more (Allport, 1954), such that workers in occupations with more teamwork might be less affected by a change in the racial composition of their peers. Results are presented in Panel A of Table 3. We find that the negative effects of a non-white vs. white death on the retention of non-white incumbents are only driven by workers in white-collar occupations, while we do not find significant effects on blue-collar workers. Moreover, the effects appear to be weaker in jobs that involve more teamwork, suggesting that in these jobs, coworker interactions may have reduced preferences for same-race coworkers. Overall, the results highlight that changes in the racial coworker composition may have different effects across different segments of the labor market.

Tenure of incumbent and deceased worker. We further examine heterogeneity in the job experience that the incumbent and the deceased worker have in common. In particular, we consider the minimum value of the incumbent's and the deceased worker's tenure at the firm.

[18] For this, we map the profession-based CBO classification in RAIS to the skill-oriented *International Standard Classification of Occupations* (ISCO-88) using the crosswalk provided by Muendler *et al.* (2004). This allows us to distinguish between white- and blue-collar occupations based on the ten major ISCO-88 occupations.

[19] We match the CBO classification to O*NET jobs using crosswalk tables from the U.S. Bureau of Labor Statistics (2023). High teamwork occupations are then defined as those with above-median ratings in the 'Work with work group or team' category.

This measure captures variation in the amount of time that both have spent together in the peer group and thus reflects the incumbents' opportunity to develop a stronger personal connection with the deceased worker. In Panel B of Table 3, we split the sample into quartiles of the minimum tenure value. The negative retention effects are largest and only significant in the bottom two quartiles, i.e. where the incumbent has less tenure alongside the deceased worker in the firm. The results suggest that a stronger personal relationship with the deceased worker does not explain the increase in separations. Instead, the results are consistent with the effect heterogeneity observed for occupations that require more teamwork. As coworkers interact more with each other, their ties become less racially segregated, and, thereby, they are less likely to respond to whether the departing worker is of the same or a different race.

Initial racial composition. We also investigate whether the effects differ along the racial composition of the peer group before the death. In Panel C of Table 3, we split our sample by the initial non-white share in the peer group using bins of 25 percentage points. We only find significantly negative retention effects for non-white workers in peer groups with a high initial non-white share. These results suggest that non-white workers who are initially in the majority react most strongly to the loss of a same-race coworker, while when non-white workers are in the minority, a further decline in the non-white share does not significantly affect their retention. However, we emphasize that because naturally, we observe smaller samples of non-white incumbents in peer groups with lower initial non-white shares, we, unfortunately, lack statistical power to rule out non-negligible effect sizes in these peer groups. For example, the coefficient for the 0-25% non-white share is nearly identical to that of the 50%-75% group, while only the latter is significant. Still, smaller effect sizes for peer groups in which non-white workers are in the minority are again consistent with the contact hypothesis. In these peer groups, non-whites have more interactions with white coworkers which may have reduced racial bias in their coworker preferences. Alternatively, the results can also be explained by models of labor market sorting based on preferences for job amenities (Rosen, 1986)[20]: Non-white workers who value working with similar peers less strongly may sort into jobs with smaller proportions of non-white coworkers and also respond less strongly to an unexpected decrease in the share of same-race coworkers.[21]

[20] Maestas *et al.* (2023) and Nagler *et al.* (2023) provide evidence that workers with higher valuations of different working conditions tend to sort into jobs that provide more of those amenities.

[21] The results are not in line with theories of tokenism that would predict larger effects for minority groups that suffer greater social isolation (Turco, 2010).

4.4 Post-Separation Transitions

To further understand the potential drivers behind the higher separation rates of non-white workers following the loss of a non-white peer, we use the matched employer-employee data to track separated workers to their new jobs. For 57.4% of the incumbent workers (both non-white and white) who separate from their current job, we observe a new formal job in the same or the following calendar year. For these workers, we can compare the characteristics of the new and the old job. To investigate whether workers tend to switch to jobs with a higher non-white share, we re-estimate model (1) using the difference between the non-white share in the new job (at the time of joining) and the non-white share in the old job (at the time of separation) as dependent variable.[22]

The results, presented in column (1) of Table 4, show that the loss of a non-white vs. white worker in the old job leads non-white workers to transition to a new job with a non-white share that is 3.4 percentage points higher than in the old job. In other words, they counteract the initial decline in the non-white coworker share by moving again to peer groups with a higher non-white share. These results align with the hypothesis that preferences for same-race coworkers shape the initial separations of non-white incumbents. Interestingly, we also find a significant positive effect on the change in the non-white share for white incumbents. Considering that these workers separated from their old jobs after experiencing the loss of a non-white vs. white worker, it is likely that we are capturing a subset of white workers who may have a preference for greater racial diversity and who thus also move to jobs with a higher non-white share.

In column (2) of Table 4, we also test whether transitioning workers are willing to accept wage penalties to work again in teams with a higher non-white share. The results show negative but small (and only for white incumbents marginally significant) effects on the change in log wages between the new and the old jobs. Thus, we do not find strong evidence for compensating wage differentials of working with more non-white peers.

4.5 Interaction with Gender Composition

So far, we have seen that a decrease in the non-white share of the peer group negatively affects the retention of non-white incumbents. While we consider racial groups as a primary delineation that may drive heterogeneity in workplace peer effects, other peer characteristics – such as gender – may also influence employee retention decisions. Much of the workplace demography literature considers both race and gender as important factors in determining the effects of coworker composition (Elvira and Cohen, 2001; Hirsch *et al.*, 2020; Linos *et al.*, 2024). In Brazil,

[22] The racial composition is measured, as before, among all employees working in the same 4-digit occupation as the focal worker.

occupations are segregated even more by gender than by race (Salardi, 2016), and productivity peer effects are found to be larger between same-gender than between cross-gender coworkers (Messina *et al.*, 2023).

Main effects of gender. We estimate the retention effects of the gender coworker composition by exploiting deaths of female and male workers. For this, we adjust model (1) as follows:

$$\begin{aligned} Y^g_{ijt} = \alpha^g + \beta^g_1 \mathbb{1}[\,\text{Deceased} = \text{female}]_j \\ + \sum_m \sum_n \gamma^g_{mn} \mathbb{1}[\#\text{F} = m, \#\text{M} = n] + \delta^g X_{ij} + \epsilon^g_{ijt}. \end{aligned} \tag{3}$$

The regressor of interest is now a dummy that captures whether the deceased worker is female. Importantly, we control for the number of female workers m and male workers n in the peer group before the death. Thus, we assume that when comparing peer groups with exactly the same initial number of female and male workers, the gender of the deceased worker is exogenous to all other forces that drive the retention of incumbents. Additional control variables in X_{ij} are the same as in model (1). We focus on the average effects across post-death event years by pooling the estimation for all $t = \{0, 1, 2, 3\}$.

Results are shown in Table 5, separately for female incumbents (column (1)) and male incumbents (column (4)). For both groups, we find precisely estimated zero effects. These results suggest that a sudden change in the gender composition of the peer group does not affect the retention of employees, in contrast to our results for the racial composition of the peer group.

Interaction between gender and race. We also explore the potential interaction between the gender and the racial composition in driving incumbents' retention. The intersectionality literature has documented the double disadvantage that non-white women face in the labor market (Crenshaw, 1989; Fernandes, 2015; Purdie-Vaughns and Eibach, 2008; Smith *et al.*, 2019). Thus, it is of particular interest whether non-white women and men are affected differently by losing coworkers of different races and genders. To analyze in a more granular way what peer characteristics affect retention decisions, we estimate the following model:

$$\begin{aligned} Y^{gr}_{ijt} = \alpha^{gr} + \beta^{gr}_1 \mathbb{1}[\,\text{Deceased} = \text{female}]_j + \beta^{gr}_2 \mathbb{1}[\,\text{Deceased} = \text{non-white}]_j \\ + \beta^{gr}_3 \mathbb{1}[\,\text{Deceased} = \text{female \& non-white}]_j \\ + \sum_k \sum_l \sum_m \sum_n \gamma^{gr}_{klmn} \mathbb{1}[\#\text{NW} = k, \#\text{W} = l, \#\text{F} = m, \#\text{M} = n] + \delta^{gr} X_{ij} + \epsilon^{gr}_{ijt}, \end{aligned} \tag{4}$$

The model combines (1) and (3), estimating the joint effects of the race and gender of the deceased worker, as well as its interaction. We control for the initial racial and gender composition by adding interacted dummies for the exact number of non-white, white, female, and male workers in the peer group before the death.

Table 5 shows results separately for non-white female, white female, non-white male, and white male incumbents. We observe the strongest effects for non-white females. For them, the death of a female coworker lowers the likelihood of staying at the firm by 1.9 percentage points, and the death of a non-white coworker reduces retention by even 2.5 percentage points. However, the interaction of the deceased worker being female and non-white is positive and almost offsets the main effects. Adding up all coefficients, the loss of a female non-white worker does not significantly affect the retention of female non-white incumbents. They are only affected by losing a female white or a male non-white coworker (all relative to our omitted category: the loss of a male white worker). Apart from female non-white incumbents, we also find that male non-white incumbents are more likely to leave the firm after the death of a non-white coworker, but for them, the gender of the deceased coworker does not matter. Moreover, for white (female and male) incumbents we do not see any effects.

These results align very closely with the findings from Linos *et al.* (2024) for new hires in a professional services firm in the United States. There, the only race and gender group whose turnover is affected by the share of white coworkers is Black women. Moreover, it is mainly the share of white male coworkers, not of white female coworkers, that is found to increase the turnover of Black women. This is consistent with our result that only the death of a non-white male, instead of a non-white female, affects the retention of non-white females. Finally, finding larger effects for female rather than male incumbents is also in line with Elvira and Cohen (2001) who find that a higher proportion of women leads to a significant reduction of turnover among women, whereas an increase of men does not cause a decrease in quitting among men.

5 Conclusion

We study the impact of having same-race coworkers on workers' retention at the firm. Using matched employer-employee data from Brazil, we exploit the deaths of workers from different racial groups as an exogenous shock to the racial composition of peer groups. We find that the death of a non-white worker – relative to the death of a white worker – decreases the likelihood of staying at the same firm for non-white incumbents. In contrast, we find little evidence that the racial coworker composition matters for white incumbents. The death-induced decline in the non-white share of the peer group makes non-white incumbents more likely to quit their current job – while not affecting their likelihood of being laid off – and to move to a new job where the non-white share is again higher than in their old job. Together, these results suggest that non-white workers' homophilic preferences for working with same-race peers explain the reduced retention rates.

Heterogeneity analyses unveil that the effects are concentrated in white-collar jobs that require little teamwork, among incumbents who have low common firm tenure alongside the deceased worker, and in peer groups in which non-white workers were initially in the majority. The results suggest that racial coworker preferences may be particularly large in work environments with few intergroup interactions among coworkers. Finally, we find that shocks to the gender composition appear less important in driving retention decisions than changes in the racial composition. At the same time, the effects of having same-race peers differ across gender groups with non-white women being most strongly affected by their peer group composition. Overall, our findings highlight the presence of large heterogeneities and asymmetries in the effects of the racial peer group composition. Studying only a small set of firms or bundling together different race and gender groups may thus not give us a complete picture of peer effects at work.

Our results underscore the importance of post-hiring dynamics in driving differential labor market outcomes of non-white and white workers. Coworker preferences can create peer effects that contribute to high levels of racial segregation across workplaces in Brazil. Therefore, public policies and firms aiming for greater diversity should not only focus on eliminating discrimination in hiring, but also pay attention to coworker interactions after hiring. One should be cautious in concluding from our results that more segregated workplaces are beneficial for the careers of non-white workers. Instead, the results call for more research on the potential role of firms' staffing and promotion policies in mitigating peer effects that exacerbate racial inequalities at work. For example, our findings indicate that retention effects are less pronounced in work environments with more coworker interactions, which suggests that fostering stronger interpersonal relationships within teams could serve as a potential approach for firms to reduce racial biases in the workplace.

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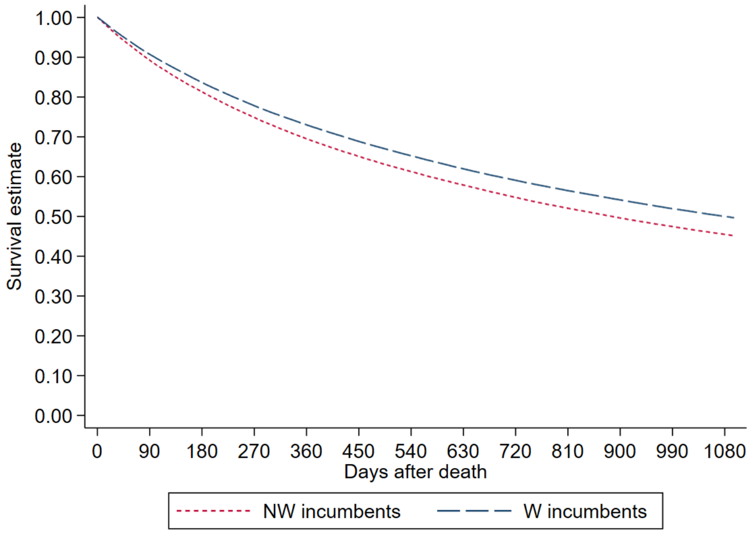
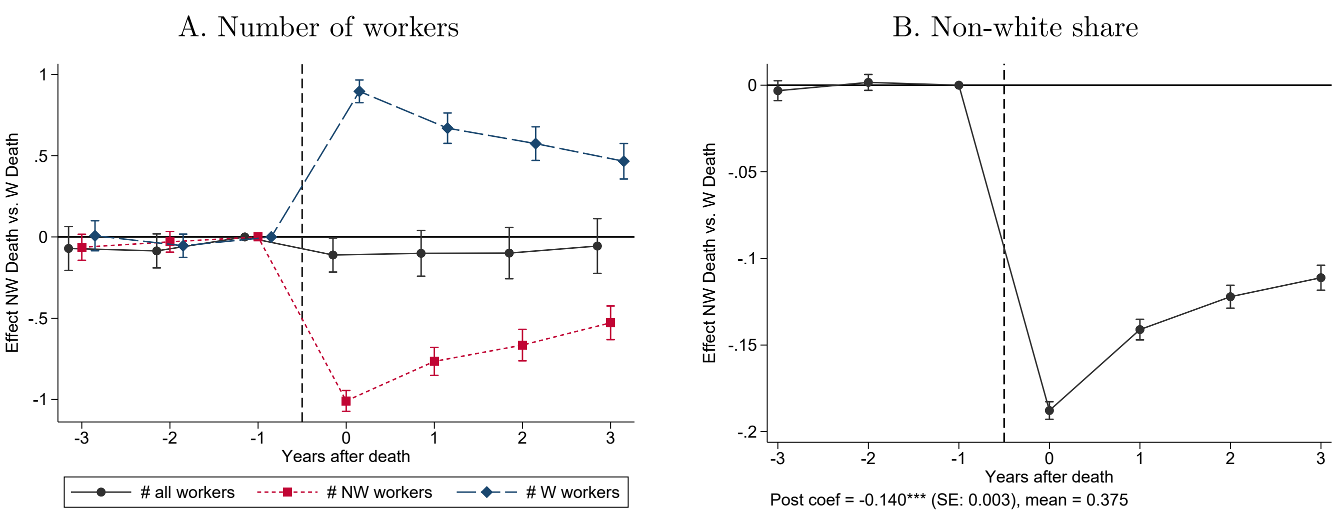
Figures and Tables

Figure 1: Kaplan-Meier estimates of incumbent retention



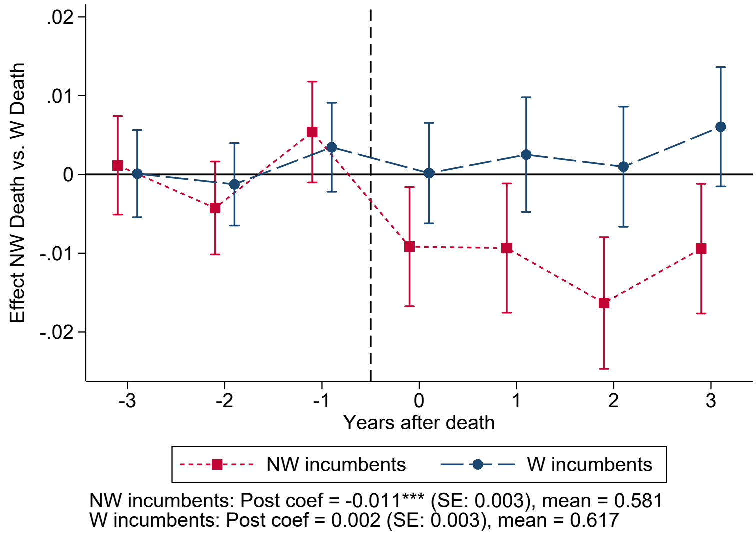
Notes: The figure shows the estimated Kaplan–Meier survivor function for non-white ($N = 156,743$) and white ($N = 256,318$) incumbent workers. The y-axis represents the probability that a worker has stayed at the firm until a given point in time. The x-axis represents the number of days since the death of the coworker.

Figure 2: Effects on size and non-white share of peer group



Notes: The figure reports the coefficients β_t estimated in model (2) for the effect of a non-white vs. white worker death on the total number of workers, the total number of non-white workers, and the total number of white workers (Panel A), as well as the non-white share in the peer group (Panel B). N = 340,732 peer group × year observations. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below Panel B, the average effect across all post-death periods and the corresponding sample mean are reported.

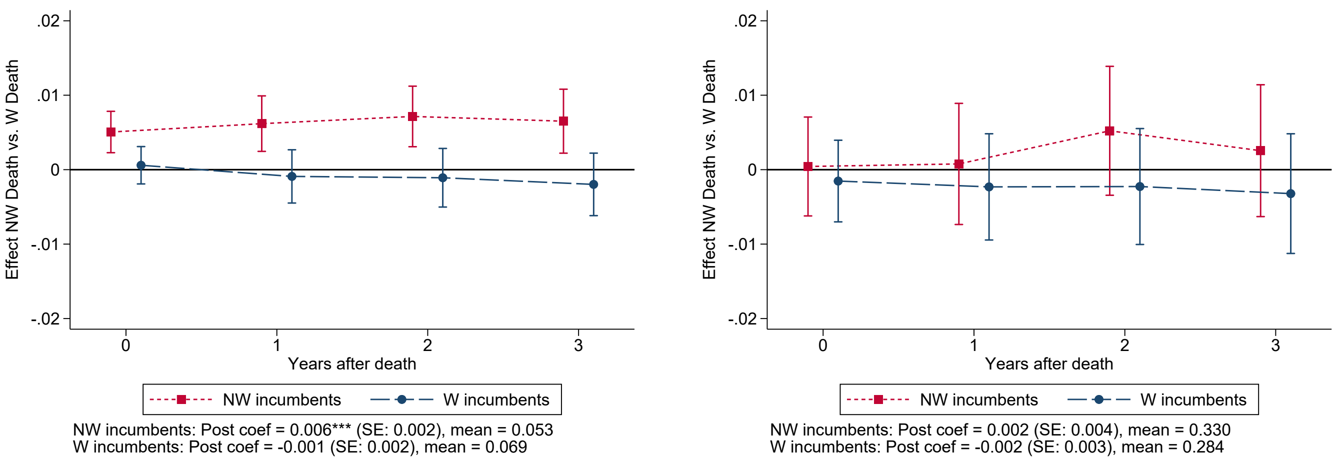
Figure 3: Effects on incumbent retention



Notes: The figure reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on incumbent workers' retention, i.e. a dummy for whether they work at the same establishment as at the time of death. The model is run separately for each event year $t = \{-3, -2, ..., 3\}$ and for incumbents of race $r = \{$non-white, white$\}$. N = 156,743 non-white incumbents and 256,318 white incumbents in each event year. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below the graph, the average effect across all post-death periods and the corresponding sample mean of retention are reported.

Figure 4: Effects on incumbent quits and layoffs

A. Quit until t B. Laid off until t



Notes: The figures report the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on a dummy for the incumbent having quit until event year t (Panel (a)) and a dummy for the incumbent being laid off until t (Panel (b)). The model is run separately for each event year $t = \{-3, -2, ..., 3\}$ and for incumbents of race $r = \{$non-white, white$\}$. N = 156,743 non-white incumbents and 256,318 white incumbents in each event year. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below each graph, the average effect across all post-death periods and the corresponding sample means are reported.

Table 1: Summary statistics

	Deceased workers			Incumbent workers		
	All	Non-white	White	All	Non-white	White
Non-white	0.36	1	0	0.38	1	0
	(0.48)			(0.49)		
Age	41.72	40.28	42.53	36.44	35.60	36.96
	(12.97)	(12.72)	(13.04)	(11.28)	(10.98)	(11.43)
Male	0.83	0.84	0.83	0.78	0.80	0.77
	(0.37)	(0.36)	(0.38)	(0.41)	(0.40)	(0.42)
Education (years)	10.17	9.84	10.36	10.40	10.07	10.60
	(3.24)	(3.21)	(3.24)	(3.17)	(3.15)	(3.17)
Monthly wage (R$, CPI 2018)	2,412	2,039	2,618	2,230	1,908	2,427
	(2,806)	(2,196)	(3,073)	(2,051)	(1,560)	(2,278)
Tenure (years)	4.94	4.41	5.23	3.81	3.41	4.06
	(5.93)	(5.39)	(6.18)	(4.96)	(4.53)	(5.18)
N	48,676	17,346	31,330	413,061	156,743	256,318

Notes: The table shows mean characteristics of deceased workers and incumbent workers in their peer groups. See Section 3.1 for the definition of deceased and incumbent workers in the RAIS data. Standard deviations are reported in parentheses. All variables are measured at the time of death.

Table 2: Balance of incumbent characteristics

	Dep. var.: Deceased is non-white					
	All incumbents		Non-white incumbents		White incumbents	
	(1)	(2)	(3)	(4)	(5)	(6)
Non-white	0.356***	-0.104***				
	(0.003)	(0.001)				
Male	0.026***	-0.003	0.050***	0.001	0.013***	-0.005
	(0.004)	(0.004)	(0.007)	(0.005)	(0.004)	(0.004)
Education (Ref.: non/elementary)						
Middle school	-0.011***	0.004	-0.007	0.006	-0.013***	0.003
	(0.004)	(0.004)	(0.006)	(0.005)	(0.004)	(0.004)
High school	0.005	0.002	0.024***	0.008	-0.009*	-0.001
	(0.005)	(0.004)	(0.007)	(0.005)	(0.005)	(0.004)
University	-0.007	0.002	-0.023	-0.005	-0.024***	0.003
	(0.009)	(0.007)	(0.016)	(0.012)	(0.009)	(0.007)
Age (Ref.: 16-25)						
26-35	0.026***	-0.002	0.034***	-0.003	0.021***	-0.001
	(0.003)	(0.002)	(0.004)	(0.003)	(0.003)	(0.003)
36-45	0.026***	-0.001	0.036***	0.001	0.020***	-0.001
	(0.004)	(0.003)	(0.005)	(0.004)	(0.004)	(0.003)
46-55	0.023***	-0.002	0.036***	0.004	0.017***	-0.005
	(0.004)	(0.003)	(0.006)	(0.005)	(0.004)	(0.004)
56-65	0.023***	-0.004	0.034***	0.001	0.019***	-0.007
	(0.005)	(0.004)	(0.009)	(0.007)	(0.006)	(0.005)
Log(wage)	-0.111***	-0.004	-0.162***	-0.001	-0.084***	-0.006
	(0.005)	(0.004)	(0.007)	(0.006)	(0.005)	(0.004)
Log(tenure)	0.004***	0.001	0.014***	0.002*	-0.003***	-0.000
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
Log(firm size)	0.029***	0.001	0.024***	0.000	0.033***	0.002
	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)
N	413,061	413,061	156,743	156,743	256,318	256,318
Pre-death #NW $\times$ #W FE	No	Yes	No	Yes	No	Yes
P-value joint signif.	0.000	0.867	0.000	0.481	0.000	0.365

Notes: The table reports results from regressions of a dummy for the deceased worker being non-white on incumbent characteristics, separately for all incumbents, non-white incumbents, and white incumbents. All regressions include year-of-death fixed effects. Columns (2), (4), and (6) add interacted fixed effects for the exact number of non-white and white workers in the peer group before the death (calculated as the sum of incumbents on the day of death plus the deceased worker). The table also reports the results of testing the joint significance of all incumbent characteristics (except 'non-white'). Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Effect heterogeneity

	(1)	(2)	(3)	(4)
Occupation type	White collar	White collar	Blue collar	Blue collar
Teamwork	High	Low	High	Low
[A.1] Non-white incumbents				
β	-0.016**	-0.025***	-0.003	-0.010
	(0.007)	(0.007)	(0.007)	(0.007)
N	131,334	126,038	174,582	169,046
Mean	0.577	0.597	0.549	0.600
[A.2] White incumbents				
β	0.001	-0.002	0.004	-0.003
	(0.007)	(0.006)	(0.006)	(0.006)
N	212,710	218,006	269,688	275,224
Mean	0.613	0.637	0.592	0.626
Common tenure of incumbent & deceased	1st quartile	2nd quartile	3rd quartile	4th quartile
[B.1]: Non-white incumbents				
β	-0.013**	-0.016***	-0.009	-0.003
	(0.006)	(0.006)	(0.006)	(0.006)
N	169,220	159,904	156,880	140,968
Mean	0.431	0.536	0.625	0.761
[B.2]: White incumbents				
β	0.005	0.004	-0.001	0.003
	(0.006)	(0.005)	(0.005)	(0.005)
N	246,784	250,636	256,196	271,656
Mean	0.462	0.565	0.647	0.778
Initial NW share	[0-25)	[25-50)	[50-75)	[75-100]
[C.1]: Non-white incumbents				
β	-0.010	-0.006	-0.011**	-0.018**
	(0.012)	(0.006)	(0.005)	(0.007)
N	57,056	116,704	161,964	291,248
Mean	0.575	0.579	0.586	0.579
[C.2]: White incumbents				
β	-0.006	0.006	0.008	0.025*
	(0.005)	(0.005)	(0.006)	(0.015)
N	668,708	217,416	106,064	33,084
Mean	0.630	0.605	0.587	0.545

Notes: The table reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on incumbent workers' retention. We report results averaged across all post-death event periods $t = \{0, 1, 2, 3\}$ and separately for incumbents of race $r = \{\text{non-white, white}\}$. Panel A divides the sample by white-collar and blue-collar occupations that involve above vs. below median levels of teamwork, Panel B by quartiles of the minimum of the incumbent's and the deceased worker's tenure , and Panel C by the inital non-white share in the peer group (see Section 4.3 for details). Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Post-separation transitions

Dep. var.:	Diff. NW share (1)	Diff. log wages (2)
[A] Non-white incumbents		
β	0.034***	-0.004
	(0.007)	(0.007)
N	55,400	55,314
Mean	-0.019	0.110
[B] White incumbents		
β	0.027***	-0.012*
	(0.005)	(0.007)
N	85,828	85,720
Mean	0.018	0.103

Notes: The table reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on the change in job characteristics of incumbent workers who move to a new job. The sample consists of all incumbent workers who separate from their initial job and start a new job in the same or the following calendar year. The dependent variables are the changes in the non-white share of the peer group (column 1) and log wages (column 2) between the new job (at the time of joining) and the old job (at the time of separation). Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$ *** $p < 0.01$

Table 5: Interaction between gender and race composition

Gender of incumbent	Female			Male		
Race of incumbent	All	Non-white	White	All	Non-white	White
	(1)	(2)	(3)	(4)	(5)	(6)
Deceased = female	-0.000	-0.019*	0.002	-0.001	-0.005	-0.002
	(0.004)	(0.011)	(0.006)	(0.005)	(0.010)	(0.007)
Deceased = non-white		-0.025***	-0.011		-0.010**	0.006
		(0.010)	(0.008)		(0.004)	(0.004)
Deceased = female & non-white		0.038***	0.009		0.005	0.009
		(0.014)	(0.011)		(0.014)	(0.013)
N	358,736	123,620	235,116	1,293,508	503,352	790,156
Mean	0.595	0.579	0.604	0.606	0.581	0.621

Notes: The table shows results on how incumbent workers' retention is affected by the gender and race of the deceased worker. Columns (1) and (4) report the coefficients β_1^g estimated in model (3) for the effect of a female vs. male death, separately for incumbents of gender $g = \{\text{female, male}\}$. Columns (2), (3), (5), and (6) report coefficients β_1^{gr}, β_2^{gr}, and β_3^{gr} estimated in model (4) for the joint effects of the deceased workers' gender and race, separately for incumbents of gender $g = \{\text{female, male}\}$ and race $r = \{\text{non-white, white}\}$. All coefficients refer to the average effect across all post-death event periods $t = \{0, 1, 2, 3\}$. Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$ *** $p < 0.01$

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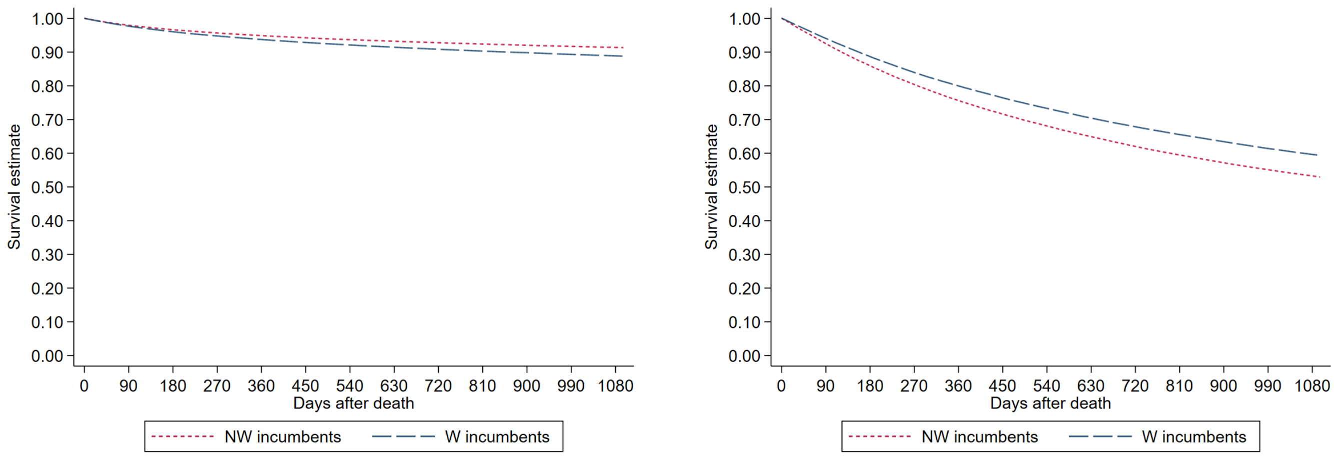
Figure A.1: Kaplan-Meier estimates of incumbent retention by quits and layoffs

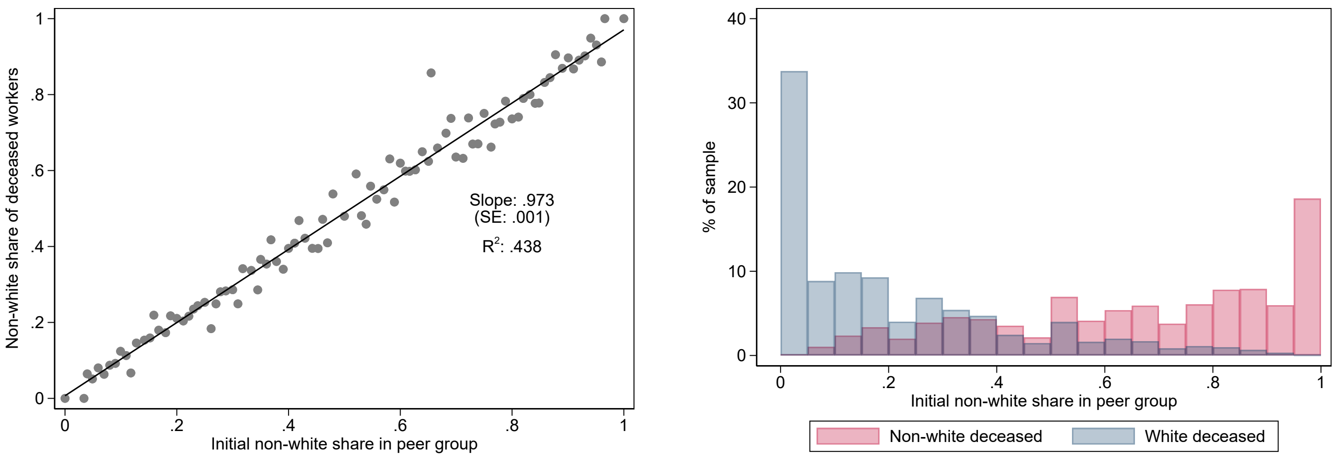
A. Quit B. Layoffs



Notes: The figure shows the estimated Kaplan–Meier survivor function for non-white ($N = 156,743$) and white ($N = 256,318$) incumbent workers, separately for those who quit and those who were laid off. The y-axis represents the probability that a worker has stayed at the firm until a given point in time. The x-axis represents the number of days since the death of the coworker.

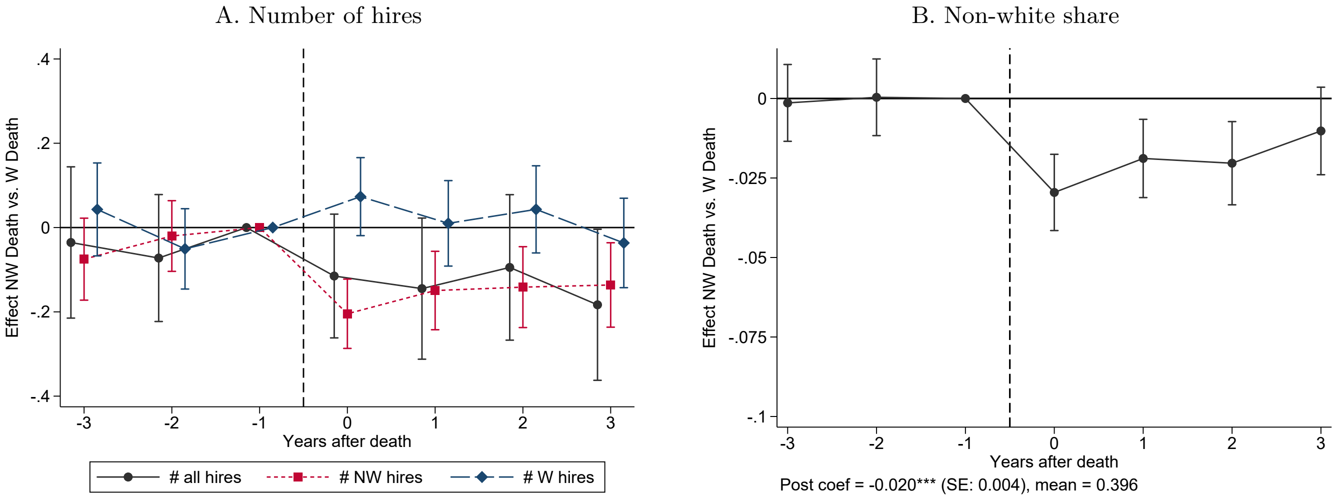
Figure A.2: Initial non-white share and race of deceased worker

A. Relation B. Distribution



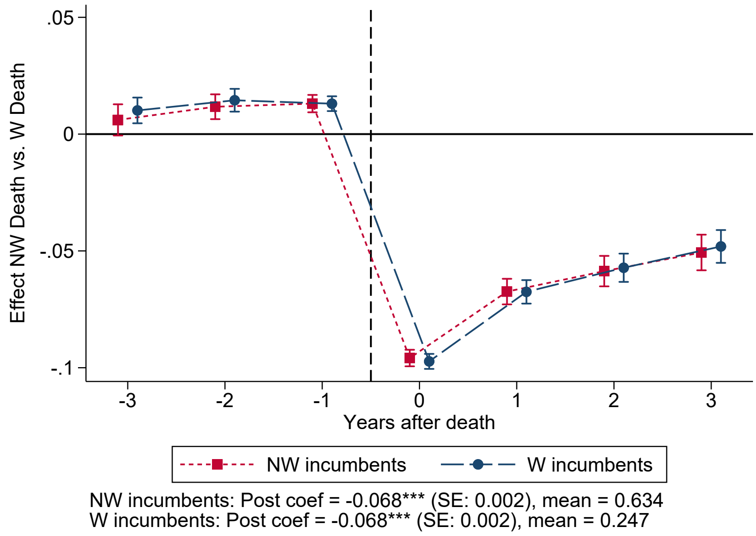
Notes: Figure A shows the relationship between the initial non-white share among all workers in the peer group (all incumbents on the day of death plus the deceased worker) and the race of the deceased worker. The dots report the non-white share of deceased workers for each one-percentage-point bin of the initial non-white share. Figure B shows the distribution of the initial non-white share for peer groups with a non-white vs. white deceased worker. Peer groups are always weighted by their size before the death.

Figure A.3: Effects on number and non-white share of hires



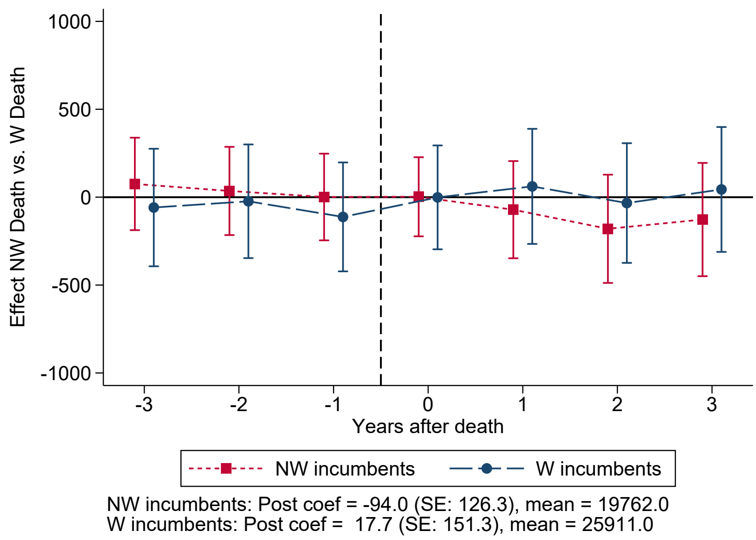
Notes: The figure reports the coefficients β_t estimated in model (2) for the effect of a non-white vs. white worker death on the number of hires, the number of non-white hires, and the number of white hires (Panel A), as well as the non-white share of hires in the peer group (Panel B). N = 340,732 peer group $\times$ year observations. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below Panel B, the average effect across all post-death periods and the corresponding sample mean are reported.

Figure A.4: Effects on non-white share of incumbents' coworkers



Notes: The figure reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on the (leave-out) non-white share of incumbent workers' coworkers in their initial peer group. The model is run separately for each event year $t = \{-3, -2, ..., 3\}$ and for incumbents of race r = {non-white, white}. N = 156,743 non-white incumbents and 256,318 white incumbents in each event year. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below the graph, the average effect across all post-death periods and the corresponding sample means are reported.

Figure A.5: Effects on earnings



Notes: The figure reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on incumbent workers' yearly real earnings. The model is run separately for each event year $t = \{-3, -2, ..., 3\}$ and for incumbents of race $r = \{\text{non-white, white}\}$. $N = 156{,}743$ non-white incumbents and 256,318 white incumbents in each event year. 95% confidence intervals are depicted for standard errors clustered at the peer group level. Below the graph, the average effect across all post-death periods and the corresponding sample means are reported.

Table A.1: Balance of local mortality rates

Dep. var.:	Deceased is non-white	
	(1)	(2)
# of deaths	0.003***	0.001
(per 1000 population)	(0.001)	(0.001)
NW share of deaths	0.831***	-0.016
	(0.019)	(0.017)
# of deaths through homicide	0.004	0.005
(per 1000 population)	(0.041)	(0.033)
NW share of deaths through homicide	0.128***	0.028
	(0.023)	(0.019)
# of deaths through firearms	0.105**	-0.009
(per 1000 population)	(0.046)	(0.037)
NW share of deaths through firearms	-0.022	-0.021
	(0.019)	(0.015)
N	44,711	44,711
Pre-death #NW $\times$ #W FE	No	Yes
P-value joint signif.	0.000	0.404

Notes: The table reports results from peer-group level estimations that regress a dummy for the deceased worker being non-white on local mortality rates measured in the year of death. In both columns, the regression includes year-of-death fixed effects. Column (2) adds interacted fixed effects for the exact number of non-white and white workers in the peer group before the death (calculated as the sum of incumbents on the day of death plus the deceased worker). The table also reports the results of testing the joint significance of all local mortality variables. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.2: Robustness: control variables

Dep. var.:	Retention (1)	Quit (2)	Laid off (3)	Earnings (4)
[A] Incumbent controls + state × 1dgt occ × 1dgt ind FE (baseline)				
[A.1] Non-white incumbents				
β	-0.011***	0.006***	0.002	-94.0
	(0.003)	(0.002)	(0.004)	(126.3)
[A.2] White incumbents				
β	0.002	-0.001	-0.002	17.7
	(0.003)	(0.002)	(0.003)	(151.3)
[B] Baseline - incumbent controls				
[B.1] Non-white incumbents				
β	-0.010**	0.006***	0.002	-66.6
	(0.004)	(0.002)	(0.004)	(249.0)
[B.2] White incumbents				
β	-0.000	0.000	-0.001	-379.2
	(0.004)	(0.002)	(0.003)	(304.3)
[C] Baseline - state × 1dgt occ × 1dgt ind FE				
[C.1] Non-white incumbents				
β	-0.008**	0.006***	-0.000	-121.4
	(0.004)	(0.002)	(0.004)	(138.8)
[C.2] White incumbents				
β	0.004	-0.001	-0.004	-53.2
	(0.003)	(0.002)	(0.003)	(172.3)
[D] Baseline + state × 2dgt occ × 2dgt ind FE				
[D.1] Non-white incumbents				
β	-0.011***	0.007***	0.001	-157.9
	(0.004)	(0.002)	(0.004)	(110.5)
[D.2] White incumbents				
β	0.002	-0.000	-0.003	-93.9
	(0.003)	(0.002)	(0.003)	(133.1)
[E] Baseline + local mortality controls				
[E.1] Non-white incumbents				
β	-0.012***	0.006***	0.003	-124.7
	(0.003)	(0.002)	(0.004)	(129.7)
[E.2] White incumbents				
β	0.001	-0.000	-0.002	7.2
	(0.003)	(0.002)	(0.003)	(157.1)

Notes: The table reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on incumbent workers' retention, quits, layoffs, and earnings when including different sets of control variables. In all panels, we include fixed effects for the initial number of non-white and white workers in the peer group and for the year of death. 'Incumbent controls' and 'local mortality controls' refer to the variables shown in Tables 2 and A.1, respectively. We report results averaged across all post-death event periods $t = \{0, 1, 2, 3\}$ and separately for incumbents of race $r = \{\text{non-white, white}\}$. $N =$ 626,972 for non-white incumbents and $N =$ 1,025,272 for white incumbents. Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.3: Robustness: interaction between race of deceased and incumbent workers

Dep. var.:	Retention (1)	Quit (2)	Laid off (3)	Earnings (4)
[A] Baseline				
$\mathbb{1}$[Deceased = non-white]$_j$	-0.002	-0.001	-0.002	17.7
	(0.003)	(0.002)	(0.003)	(151.4)
$\mathbb{1}$[Deceased = non-white]$_j$	-0.013***	0.007***	0.005	-111.7
$\times\mathbb{1}$[Incumbent = non-white]$_i$	(0.004)	(0.002)	(0.004)	(155.3)
[B] Add peer group fixed effects				
$\mathbb{1}$[Deceased = non-white]$_j$	-0.014***	0.006***	0.006	-39.1
$\times\mathbb{1}$[Incumbent = non-white]$_i$	(0.004)	(0.002)	(0.004)	(111.4)

Notes: The table reports results from a version of model (1) that is estimated jointly for non-white and white incumbents. Panel A includes interactions between all explanatory variables specified in model (1) and a dummy for whether the incumbent is non-white. Panel B additionally includes peer group fixed effects and only identifies the differential effect of a non-white vs. white worker death on non-white vs. white incumbents. We report results averaged across all post-death event periods $t = \{0, 1, 2, 3\}$. N = 1,652,244 incumbents. Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.4: Robustness: cox regression

Dep. var.:	Separated (1)	Quit (2)	Laid off (3)
[A] Non-white incumbents			
β	0.021***	0.008***	0.004
	(0.007)	(0.002)	(0.006)
[B] White incumbents			
β	-0.005	-0.001	-0.004
	(0.006)	(0.002)	(0.005)

Notes: The table reports results from a Cox proportional hazard model for the effect of a non-white vs. white worker death on incumbent workers' separations. The model is run separately for non-white incumbents (Panel A, N = 156,743) and white incumbents (Panel B, N = 256,318), as well as for any separations, quits, and layoffs. A positive coefficient indicates a higher likelihood of leaving the current job. Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.5: Robustness: peer group definition

Dep. var.:	Retention (1)	Quit (2)	Laid off (3)	Earnings (4)
[A] Maximum 30 incumbents in same 4dgt occupation (baseline)				
[A.1] Non-white incumbents ($N = 626,972$)				
β	-0.011***	0.006***	0.002	-94.0
	(0.003)	(0.002)	(0.004)	(126.3)
[A.2] White incumbents ($N = 1,025,272$)				
β	0.002	-0.001	-0.002	17.7
	(0.003)	(0.002)	(0.003)	(151.3)
[B] Maximum 30 incumbents in same 6dgt occupation				
[B.1] Non-white incumbents ($N = 556,548$)				
β	-0.010***	0.006***	0.000	-77.1
	(0.004)	(0.002)	(0.004)	(135.5)
[B.2] White incumbents ($N = 903,704$)				
β	0.003	-0.002	-0.003	-60.9
	(0.003)	(0.002)	(0.003)	(154.0)
[C] Maximum 20 incumbents in same 4dgt occupation				
C.1] Non-white incumbents ($N = 441,280$)				
β	-0.009**	0.006***	-0.001	44.3
	(0.004)	(0.002)	(0.004)	(130.4)
[C.2] White incumbents ($N = 744,628$)				
β	0.005	-0.001	-0.006*	-2.3
	(0.003)	(0.002)	(0.004)	(153.5)
[D]Maximum 10 incumbents in same 4dgt occupation				
[D.1] Non-white incumbents ($N = 190,600$)				
β	-0.011**	0.006**	0.001	5.3
	(0.005)	(0.003)	(0.006)	(181.0)
[D.2] White incumbents ($N = 343,028$)				
β	0.011**	-0.001	-0.012**	43.2
	(0.005)	(0.003)	(0.005)	(184.7)

Notes: The table reports the coefficients β_t^r estimated in model (1) for the effect of a non-white vs. white worker death on incumbent workers' retention, quits, layoffs, and earnings when varying the definition of the peer group. Panel A reports results for incumbents working in the same 4dgt occupation as the deceased worker and a maximum of 30 incumbents in the peer group (our baseline specification). In Panel B, we restrict to incumbents in the same 6dgt occupation. In Panels C and D, we reduce to maximum peer group size to 20 and 10 incumbents, respectively. We report results averaged across all post-death event periods $t = \{0, 1, 2, 3\}$ and separately for incumbents of race $r = \{\text{non-white, white}\}$. Standard errors clustered at the peer group level are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$