Black/White Gaps for Men and Women in Employer-Sponsored Health Insurance

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Abstract: I use data from the Current Population Survey spanning the years 1988 through 2017 to analyze black/white gaps in the probability of employer-sponsored health insurance from one's own employer for male and female workers. I find that black male workers are about six percentage points less likely than white male workers to have such coverage, while black female workers are about six percentage points *more* likely than white female workers to do so. These differences persist after controlling for education. The lower rate of coverage for black men compared with white men is largely the result of lower rates of health insurance offering at the firm level, rather than eligibility conditional on offering or takeup conditional on eligibility. The higher rate of coverage for black women compared with white women is largely the result of higher takeup conditional on eligibility and, to a lesser extent, higher rates of eligibility, with very little difference in health insurance offering at the firm level. These results highlight the importance of considering men and women separately when analyzing racial differentials in the labor market.

1. Introduction

Non-wage benefits make up thirty percent of total compensation on average, with health insurance - the largest single fringe benefit – accounting for eight percent of the total (U.S. Department of Labor 2021). Economists have studied differences in wages and earnings between black and white workers for decades, but very little research has focused on racial disparities in non-wage benefits.

In this paper, I estimate black/white gaps for male and female workers in the probability of having health insurance as a fringe benefit of employment. Using Current Population Survey data spanning the thirty-year period from 1988 through 2017, I find that black male workers are about six percentage points less likely than white male workers to have such coverage, while black female workers are about six percentage points *more* likely than white female workers to do so. These differences persist after controlling for education. The lower rate of coverage for black men compared with white men is largely the result of lower rates of health insurance offering at the firm level, rather than eligibility conditional on offering or takeup conditional on eligibility. The higher rate of coverage for black women compared with white women is largely the result of higher takeup conditional on eligibility and, to a lesser extent, higher rates of eligibility, with very little difference in health insurance offering at the firm level. These results highlight the importance of considering men and women separately when analyzing racial differentials in the labor market.

2. Background

An extensive literature in economics documents persistent black/white differences in wages and/or earnings (see Altonji and Blank 1999 for a review; and more recently, Bayer and Charles 2018). As noted by Altonji and Blank (1999), much less research has examined black/white differences in non-wage compensation. Nonetheless, such papers do exist. Some examine racial differences in pensions (for example, Lazear and Rosen 1987). Focusing on employer-provided health insurance as an outcome, a number of papers document that on average, black workers are less likely than white workers to have their own employer-provided health insurance (Kalleberg, Reskin, and Hudson 2000; Cubbins and Parmer 2001; White-Means and Hersch 2005; Levy 2006) while others report blacks workers are more likely to have such coverage (Cooper and Schone 1997; Haas and Swartz 2007).

Some papers in this literature further distinguish between black/white gaps in health insurance for women and those for men. All of these studies find that black male workers are less likely than white male workers to have their own employer-sponsored coverage, while black female workers are *more* likely than white female workers to do so (Herch and White-Means 1993; Meyer and Pavalko 1996 [analyzing women only]; Monheit and Vistnes 2000; Dushi and Honig 2005; Keene and Prokos 2007; Semyonov et al. 2010; Kristal et al. 2018).[1] However, the relative advantage of black women compared to white women in this particular component of compensation has, as far as I can tell, been largely unnoticed in the wider literature. I speculate that this is for three reasons. First, labor economists studying wage inequality are far more likely to focus on men. Second, particularly during a period of intense national debate over the expansion of health insurance coverage more generally, health economists studying health insurance disparities have been more likely to focus on the entire population, not just workers,

[1] I believe the paper by Semyonov et al. (2010) contains an error: specifically, that it uses a variable in Table 1 reflecting employer-sponsored health insurance coverage from any source (including as a dependent on a spouse's policy) rather than the worker's own-employer coverage as the text implies, so that the reported rates of coverage in Table 1 are approximately equal for black and white workers of each gender. However, the text is unambiguous that white male workers and black female workers have the highest rates of coverage through their employer, and the numbers reported in the text do not match the numbers in Table 1. At least one other unpublished report from this era makes a similar error, reporting results for *any* employer-sponsored coverage as if they were rates of *own-employer* coverage.

and employer coverage from any source, not just as a fringe benefit of one's own employment; in this broader context, black women do indeed have lower rates of overall coverage than white women. Third, the focus of many of these studies is on statistically "explaining" racial differences as a function of other characteristics. With the addition of enough covariates – in particular, firm size and sector of employment – these differentials are sometimes rendered insignificant, although this begs the question of why black men and women are more likely to work in the public sector and in larger firms than are white men and women.[2] In any case, it is time to revisit black/white gaps in employer-health insurance among workers, in the context of inequalities in compensation.

3. Empirical approach

In order to shed some light on the differences in workers' own-employer coverage by race, I begin by documenting simple trends over time in own-employer sponsored health insurance coverage for groups of workers defined by race (black or white) and sex (male or female) and the black/white gaps in coverage that these trends imply. I also estimate a version of the gap adjusting for education by estimating simple linear probability models in each year, separately for male and female workers, with the following specification:

$$\Pr(Own\ EHI) = \beta_0 + \beta_1 \cdot (BLACK) + \beta_2 \cdot (education) + \varepsilon \qquad (1)$$

Education is a vector of dummies reflecting categories educational attainment (less than high school, equal to high school, some college, college or more) and is intended to proxy for skill level. In practice, the statistical "explanation" offered by controlling for education is problematic, because differences in education across groups may be the result of historical or

[2] As noted by Brown (1982), the racial discrimination provisions related to employment in the Civil Rights Act of 1964 do not apply to small firms, initially defined as private firms with fewer than 100 employees and lowered to 25 or 15 employees over time. Currently, these protections do not apply to firms with fewer than 15 employees (https://www.eeoc.gov/employers/small-business/small-business-requirements).

ongoing racism. This general problem – that controlling for covariates in regressions intended to explain disparities may explain away the very thing we want to understand - has received less attention in the economics literature than it deserves. I present both unadjusted and education-adjusted black-white gaps in employer sponsored insurance.

Next, I apply a simple algebraic decomposition to the unadjusted gaps that breaks down the gap in health insurance coverage between two groups into components attributable to health insurance offering by a worker's employer; the worker's individual eligibility for such coverage, conditional on being in a firm that offers health insurance to at least some workers; and the worker's decision to enroll in coverage, conditional on being eligible (Farber and Levy 2000; Levy 2006). That is, the rate of own-employer coverage for group *i* (C_i) is the product of the offer rate (O_i), the eligibility rate conditional on offering (E_i), and the enrollment rate conditional on eligibility, also referred to as "takeup" (T_i):

$$C_i = O_i \cdot E_i \cdot T_i \tag{1}$$

The difference in the coverage rate between two groups (e.g. black female workers versus white female workers), after some algebraic manipulation, can be written as:

$$C_2 - C_1 = (O_2 - O_1) \cdot E_2 \cdot T_2 + (E_2 - E_1) \cdot O_2 \cdot T_2 + (T_2 - T_1) \cdot O_2 \cdot E_2 + interaction \tag{2}$$

Or, in more compact notation, using Δ to denote the difference between groups 1 and 2:

$$\Delta C = \Delta O \cdot E \cdot T + \Delta E \cdot O \cdot T + \Delta T \cdot O \cdot E + interaction\ term \tag{3}$$

where the terms without subscripts represent the average rates for group 2. The three terms on the right hand side are the percentage point differences in the coverage rates between the two groups due to differences in offering, eligibility, and takeup.

An important limitation of this analysis is that I do not address the problem of selection into employment, which may bias the results. This selection problem has been noted for nearly as long as there have been comparisons of earnings or wages across groups (see, for example, Gronau 1974). Recent methods to avoid selection bias rely on having a continuous outcome variable, coupled with the assumption that the non-employed would have fallen in the lower half of the distribution had they been employed (Bayer and Charles 2018). This is not an option with a binary outcome variable, so I simply note that this is a potential problem without having any solution to offer.

4. Data

Data for my analysis come from two different supplements to the Current Population Survey (CPS): the Annual Social and Economic Supplement (ASEC), conducted every year in March, and periodic supplements on employee benefits or contingent work that have been conducted at irregular intervals since 1979.

The CPS-ASEC has been an official source of statistics on health insurance coverage since 1987.[3] The original ASEC health insurance questions recorded health insurance coverage of all family members during the calendar year prior to the survey: did the individual have any coverage at all during the year and if so, what type(s)? Private insurance is coded as being in the respondent's own name (that is, the respondent is the policyholder rather than a dependent on the policy) or not; and whether the coverage was provided by an employer or a union. Thus, for workers in the CPS, it is possible to identify those who have employer-sponsored coverage in their own name.[4] I use data from the 1989 through 2018 ASECs, corresponding to health insurance coverage held by respondents during calendar years 1988 through 2017.[5] There are

[3] Questions about health insurance were included in the CPS-ASEC starting in 1980, but the Census Bureau did not start publishing official estimates using these data until 1987 (see U.S. Census Bureau 1987). More detail about the CPS health insurance questions is available at https://www.census.gov/programs-surveys/cps/technical-documentation/user-notes/health-insurance-user-notes/health-ins-cov-meas-asec-acs.html.
[4] A very small fraction may have coverage from a former employer, sometimes known as COBRA coverage.
[5] The fact that the CPS-ASEC reference period is the prior calendar year can create confusion about which year is which. In this paper, I consistently describe results with the year of the survey reference period, rather than the year

two significant changes to the ASEC that affect the continuity of its health insurance data during this period. First, in survey year 2000 (affecting data for the 1999 reference period), the question sequence was changed to more thoroughly capture different sources of health insurance; this appears to have resulted in an increase in the number of people counted as having health insurance (Davern et al. 2002). Second, in survey year 2014, the health insurance questions were modified to capture information on coverage at the time of the survey in addition to during the prior calendar year, as well as other modifications to the question sequence, which again increased the number of people counted as having coverage (Pascale et al. 2016).[6] Thus, breaks in the data may occur in 1999 and 2013. While my analysis does not rely on the continuity of trends in the CPS-ASEC health insurance estimates, it is worth keeping in mind that some year-to-year changes may reflect measurement changes.

Since my focus is on workers, I restrict the data to those who worked at a job or business at any time during the previous calendar year, resulting in a sample of about 68,000 to 112,000 workers in each year. I further restrict the sample by dropping workers younger than 25 or older than 54; self-employed workers; and those who report their race and ethnicity as anything other than black non-Hispanic or white non-Hispanic. Together, these restriction reduce the ASEC sample for my analysis by about 50 percent. The resulting analytic sample has about 34,000 to 60,000 workers in each year, or a total of 1.35 million observations over the entire thirty-year

in which the survey was conducted, to describe the data. So, for example, when I refer to results "from 1988," these rely on data from the March 1989 CPS-ASEC, which asked questions about health insurance coverage held during 1988..

[6] A third change, the introduction of a new data processing system in 2019, affected CPS-ASEC health insurance estimates but falls outside the period I analyze. More information on the recent changes to the health insurance questions is available at https://www.census.gov/newsroom/blogs/research-matters/2019/09/cps-asec.html. This page includes a very helpful figure showing the breaks in trend in the fraction of the population that is uninsured associated with the changes described here (Figure 5).

period from 1988 through 2017. Appendix Table A1 shows the exact sample size in each year, as well as the distribution of the sample across groups defined by race and sex.

Table 1 shows mean insurance outcomes and other characteristics in the ASEC sample by race and sex, pooling all years of data from 1989 through 2018 (calendar years 1988 through 2017). Black male workers are less likely than white male workers to have own-employer health insurance coverage (62.3 percent versus 71.8 percent), while black female workers are more likely than white female workers to have this type of coverage (59.7 percent versus 57.0 percent). For both men and women, however, white workers are 8 to 10 percentage points more likely than their black counterparts to have health insurance from any source. As expected, black workers have lower levels of education than white workers and are less likely to be married; they are more likely to work in the public sector and in large firms. In terms of labor force attachment, black female workers are more likely than white female workers to have worked full-time (35 hours or more per week) and full-year (50 or more weeks per year) in the past calendar year; for men, however, black male workers are less likely than white male workers to have worked full-time and full-year. (Recall that that workers in the CPS-ASEC include anyone who worked for pay at all in the prior calendar year.)

The CPS-ASEC is valuable for its large sample and long time series, but has no information on whether workers without their own employer-sponsored coverage work for firms that offer insurance and, if they do, whether a worker is eligible to enroll in such coverage. A sequence of CPS supplements have collected this information at irregular intervals. Supplements on employee benefits were conducted in May 1988, and April 1993;[7] supplements on contingent and alternative employment arrangements (or "contingent work") were conducted in February

[7] Employee benefits supplements were also conducted in May 1979 and May 1983, but include no data on health insurance offering or eligibility for workers without their own employer-sponsored coverage.

1995, February 1997, February 1999, February 2001, February 2005, and May 2017. All of these supplements include data on employer-sponsored health insurance coverage, including whether the individual's employer offers coverage to any workers; whether the individual is eligible for such coverage (and if not, why not); and if eligible but not enrolled, why they did not take up coverage. This allows me to code employer-sponsored health insurance offering, eligibility, and takeup for each respondent.

In all of these supplements, health insurance questions were asked of individuals who were working for pay during the week prior to the interview. In the benefits supplements (1988 and 1993) these questions were asked of a random half sample of workers, while in the contingent work supplements conducted in 1995 and later, they went to all workers. As in the CPS-ASEC, I further restrict the sample by dropping workers younger than 25 or older than 54; self-employed workers; and those who report their race and ethnicity as anything other than black non-Hispanic or white non-Hispanic. Together, these restrictions reduce the benefits supplement sample for my analysis by 46 percent. I drop another 5 percent of observations due to missing data on health insurance coverage, offering, or eligibility. The resulting analytic sample has about 15,000 to 60,000 workers in each year – about half the size of the ASEC sample - or a total of 179,424 observations over all eight supplements spanning the years 1988 through 2017. Table 2 shows mean insurance outcomes and other characteristics in the benefits and contingent work supplements, pooling data across all years. Appendix Table A2 shows the exact sample size in each year, as well as the distribution of the sample across groups defined by race and sex. The reported rates of own employer-sponsored health insurance are higher in these data than in the ASEC, which most likely reflects the fact that this sample consists of individuals who were working at the time of the survey. But the patterns of coverage by race and sex are similar to those described above for the ASEC. In particular, black male workers are less likely than white male workers to report own-employer coverage, while black female workers are more likely than white female workers to do so; but again, the overall rate by health insurance from any source is higher for whites than blacks among both male and female workers. Looking at the building blocks of employer-sponsored health insurance – offering, eligibility, and takeup – black male workers have lower rates at each step than do black male workers, while the opposite is true for female workers; black female workers are more likely than their white counterparts to work in firm that offers coverage, to be eligible for that coverage when it is offered, and to take it up when they are eligible. Thus, the preliminary evidence suggests very different dynamics for male and female workers driving black/white gaps in health insurance coverage. In terms of other characteristics (education, marital status, employment sector), patterns are similar to those discussed for the ASEC. The benefits and contingent work supplements also include data on job tenure, which is an important predictor of access to employer-sponsored insurance (Farber and Levy 2000). Black male workers have lower average tenure than white male workers, while black female workers have higher average tenure than white female workers, and are less likely to have been on the job for less than a year. These differences in tenure are small but statistically significant across racial groups.

5. Results

Trends in Own-Employer Coverage: ASEC

I begin by presenting trends in own-employer health insurance coverage for workers using data from the CPS-ASEC. Figure 1 shows the fraction of white and black workers with employer-sponsored health insurance coverage from 1988 to 2017. This fraction hovers around two-thirds, dropping gradually over the three decades. At any point in time, black workers are about 5 percentage points less likely to have health insurance from their job. This figure, which combines male and female workers, obscures dramatic differences in the black-white gap for the two groups. Figure 2 presents similar trends separately for female and male workers, which look quite different from Figure 1. For female workers, those who are black generally have *higher* rates of own-employer coverage than do whites. This difference is not always significantly different from zero; in particular, between 2008 and 2015, there is little black/white gap in this benefit for female workers. For men, in contrast, black workers are consistently much less likely than white workers to have their own employer-sponsored coverage. This gap is between five and ten percentage points throughout this 30-year period.

This result should not be entirely surprising; as discussed above, studies in the existing literature have documented these facts. And yet, given that most of those studies used only one or two years' worth of data, the sheer durability of the pattern over time is striking. Any explanation for black/white differences in own-employer health insurance must allow very different patterns for male and female workers.

Unadjusted and Adjusted Black/White Gaps in Coverage: CPS-ASEC

As discussed above, using covariates like education to "explain" difference across race may unintentionally minimize the role played by race creating in these differences. It is still helpful to know whether the gaps documented in Figure 2 – and in particular, the differences between male and female workers – are in part driven by differences in observable characteristics. I therefore estimate a set of models (as described above; see equation 1) that control for education. I then plot the coefficient on the black indicator variable (β_1) from each regression and the associated 95 percent confidence intervals. Appendix Table A3 reports these regression models for the years 1988 and 2017; the models are estimated for each year and these two years are offered as examples. These adjusted black-white gaps are then juxtaposed with the unadjusted gaps estimated from a set of linear regressions estimated with no additional covariates (the unadjusted gaps are equal to the vertical differences in each year between the lines shown in Figure 2). Figure 3 shows the unadjusted and education-adjusted black-white gaps in own-employer coverage for male workers, with capped vertical lines indicating 95 percent confidence intervals. Two things are evident from this figure: the gap may be getting gradually smaller over time, and the education adjustment reduces the gap by a few percentage points. But in all years, it remains negative and significantly different from zero. Thus, for male workers, lower rates of own-employer coverage for blacks versus whites persist over time and are not explained by differences in education.

Analogous results for female workers are displayed in Figure 4. For female workers, both unadjusted and adjusted gaps fluctuate over time with no clear trend up or down. Adjusting for black female workers' lower levels of education compared to white female workers increases the positive black/white gap in own-employer coverage. While the unadjusted black/white gap in own-employer coverage among women is not significantly different from zero in every year, the education-adjusted gap is.

At this point, it would be possible to dive deeper into "explaining" black/white differentials in coverage using additional covariates. Some of these, like education, will help account for the negative gap for men while making the positive gap for women even larger; some will not. Firm size and public sector employment, in particular, would offer a partial statistical explanation for the positive black/white gap for women. But because these covariates are themselves all the result of inherently racialized processes, I take a different approach. Specifically, I turn to the benefits and contingent work supplements, and their data on health insurance offering, takeup, and eligibility, as an alternative approach to understanding what is going on. Before moving on to these decompositions, I confirm that these data show the same general patterns of own-employer coverage by race and sex over time as the CPS-ASEC data; these results are shown in Figure 5, 6, and 7, which shows generally similar patterns as those in Figure 2, 3, and 4 (admittedly with somewhat different levels, which is likely due to the sample differences already discussed).

Trends in offering, eligibility, and takeup: CPS Benefits and Contingent Work Supplements

I begin with simple trends in health insurance offering, eligibility, and takeup by race and sex. These trends are shown for male workers in Figure 8 and female workers in Figure 9. It is immediately evident that black male workers are less likely to be in firms where health insurance is offered than are white male workers, while black female workers take up health insurance at higher rates than white female workers when it is available to them. The other trends – for example, eligibility conditional on health insurance offering – look fairly similar to the naked eye for black and white workers within each sex.

Next, I carry out the decomposition of black/white coverage gaps for each group into components due to offering, eligibility, and takeup, as in equation (3) above. Table 3 lays out the decomposition, including its component parts, for all eight years of data pooled together. The top panel of Table 3 does this for male and female workers pooled together. (Since the main lesson of the analysis so far has been the importance of analyzing male and female workers separately, my point here is to illustrate just how uninformative the pooled decomposition is.) As shown in Table 3, there is on average a very small difference in this sample – just four-tenths of a percentage point – between the own-employer coverage rate for black and white workers overall. This difference is not statistically significant (standard error of 0.003). The decomposition splits the -0.4 percentage point gap in coverage into a negative 1.3 percentage point gap due to lower rates of health insurance offering for black workers, offset by a 1 percentage point higher rate of takeup. Differences in eligibility contribute nothing to the overall coverage gap, and the remaining negative 0.1 percentage point is due to the interaction term (see equation 3), which I ignore throughout because it is always negligible.

The lower panels of Table 3 display decompositions for male and female workers. The results in Panel B show that the 6.2-percentage-point deficit in own-employer coverage for black relative to white male workers is mainly due to offering (3.6 of the 6.2 points), with additional contributions from lower eligibility and takeup (1.4 and 1.1 points respectively). For female workers, shown in Panel C, the story is quite different. The majority of black female workers' higher rate of own-employer coverage, relative to white female workers (6.4 percentage points) is explained by higher takeup (4.2 points), higher eligibility (1.7 points), and a very small contribution from a higher offer rate (0.8 points).

Table 4 presents results of similar decompositions for each year. The patterns described above for the all-years sample generally hold in each year: for men, the biggest factor contributing to lower coverage for blacks than for whites is lower offering; while for women, their higher rates of coverage are largely due to higher takeup. In both cases, all three factors usually line up in terms of sign; it is not the case, for example, that black female workers are overcoming a substantial gap in offering with much higher takeup. The results for male and female workers in Panels B and C of Table 4 are presented graphically in Figures 10 and 11, which make these patterns clear.

6. Discussion

To recap the central findings of my analysis: throughout the thirty-year period from 1988 to 2017, rates of own-employer coverage were significantly *lower* for black male workers than for white male workers, and significantly *higher* for black female workers than for white female workers. These gaps were, on average, -6.2 percentage points for men and +6.4 percentage points for women. These differences persist after controlling for level of education. For men, the gap is driven mainly by lower rates of health insurance offering, with smaller contributions from lower eligibility and takeup. For women, the gap is driven mainly by higher takeup, with a substantial contribution from eligibility and very little contribution from offering.

These results have a number of implications. The first is that measures of black/white compensation inequality that incorporate health insurance in addition to wages and salaries would very likely reduce measured inequality. For female workers, this is mathematically inevitable, given higher rates of coverage for white women than black women. For male workers, the key question is whether the black/white gap in own-employer coverage is larger or smaller than the gap in earnings. This follows from the fact that compensation inequality is the weighted sum of inequality in earnings and other non-cash forms of compensation (Levy 2006). I find that the black/white gap in the probability of own-employer coverage for men is about negative 6 percentage points, on average, and possibly smaller in more recent years. This is dwarfed by the approximately 30 percent gap in earnings between black and white men in recent years (Daly et al. 2017). At the median, health insurance accounts for 11.6 percent of total compensation while wages and salaries account for 74.0 percent.[8] If the value of health insurance is about the same for all workers who have it, adding health insurance to the equation would

[8] BLS Employer Costs for Employee Compensation, May 2021. https://www.bls.gov/ect/compensation-percentile-estimates.htm

yield an estimate of compensation inequality that is only slightly smaller than earnings inequality for men.

A second implication is that it is unwise to pool data on men and women when analyzing racial gaps in employment compensation. In the case of health insurance, the intersectionality of race and sex *is* the story – one that is easy to miss by estimating black/white gaps and male/female gaps separately.[9] Equally unwise is focusing only on men, a common practice in the labor economics literature. The most common defense offered for the men-only approach is that it is difficult to model women's labor force participation.[10] But modeling participation, or developing other methods to address more general selection into employment, is a concern for men too. Given recent methods that address this selection problem in a non-parametric way (Bayer and Charles 2018), there is little justification for excluding women from analyses of racial differences in earnings. In the case of employer-sponsored health insurance, to look only at men would be to get less than half the story.[11]

These results also raise questions. The first is why health insurance appears to be less unequally distributed than earnings. This may have something to do with the group nature of health insurance; in order to achieve effective pooling of medical spending risks, the pool should be as large as possible. For self-insured health insurance plans, including those offered by the majority of large employers, non-discrimination rules under ERISA limit the extent to which

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[9] For an example of a paper that misses this story entirely, see Levy (2006). The concept of intersectionality is due to Crenshaw (1989).

[10] Consider the following footnote, taken from a recent paper about wages in a top field journal in economics: "We focus on men for two main reasons: (i) including women during early adulthood would require us to model their fertility decisions, which is outside the scope of the present analysis, and (ii) much of the literature that has studied human capital formation to which our analysis is comparable has focused on men." In other words, "we study men because we have always studied men." Note, also, the assumption implicit in their first assertion that men's fertility need not be modeled; this may be true, but this reflects societal inequities that in turn shape labor market inequities.

[11] Indeed, some studies find that black women earn more than white women, after controlling for certain factors (for example: Murnane, Willett, and Levy 1995; Neal and Johnson 1996).

employers can offer health insurance to some workers but not others without forfeiting the pre-tax status of premium payments.

Another reason why employer-sponsored insurance may be less unequally distributed with respect to race is that both black men and women are more likely than their white counterparts to work in the public sector and for large firms, and these employers are more likely to provide health benefits. As discussed above, it may be no accident that black workers are disproportionately employed by large or public employers, but rather have arisen as a result of anti-discrimination policies in employment that only apply to, or are more actively embraced by, these firms. Thus, racial disparities in coverage are *smaller* than they would be if black and white workers had similar profiles in terms of firm size and sector of employment.[12]

This last observation raises an important unresolved point. What characteristics should we control for when estimating differentials across racial groups? Versions of this question appear in the seminal work of Blinder (1973) and Oaxaca (1973), who independently developed the standard statistical decomposition used to analyze such differentials (see Appendix A). This question is central to how economists analyze race in the labor market, and it remains unanswered (Schwabisch and Kijakazi 2021). The answer may ultimately depend on the purpose of the analysis. If the purpose is to determine whether employers are treating similar workers unequally in ways that are potentially discriminatory, conditioning on observable characteristics like education may make sense. If the purpose is to understand where blacks stand relative to whites, including the effects of deep-rooted and long-standing systemic factors underlying such

[12] Indeed, regressions based on equation (1) but with additional controls for firm size and employment sector yield small but significant and positive differentials for women (about one percentage points) and large negative differentials for men (negative ten percentage points).

differences, then controlling for mechanisms like education through which inequality may be reinforced does not make sense.

Ultimately, health insurance provides a useful test case for thinking through how fringe benefits may affect total compensation inequality. Nearly every thorny issue that might arise in the analysis of wage or earnings differentials also arises for health insurance: in particular, selection bias, the challenge of choosing covariates, and the intersectionality of race and sex. Although incorporating health insurance into models of compensation inequality is unlikely to change the overall picture, it should nonetheless help to sharpen the picture's focus.

7. References

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Appendix A: Blinder (1973) and Oaxaca (1973) on Covariates

The idea that statistically controlling for differences across individuals might explain away some of what should in fact remain unexplained is present in the earliest papers on this topic. Oaxaca (1973) analyzes in male-female wage differentials while Blinder (1973) analyzes both male-female and black-white wage differentials. From Oaxaca (1973, pp. 698-699):

> One difficulty with the present formulation of the wage equation is that it controls for what many would consider to be major sources of discrimination. By controlling for broadly defined occupation, we eliminate some of the effects of occupational barriers as sources of discrimination. As a result, we are likely to underestimate the effects of discrimination.

Oaxaca's solution to this problem is to estimate models with and without controls for occupation, industry, and class of worker. The models without these controls, which Oaxaca refers to as the "personal characteristics wage regressions," include covariates measuring labor market experience, education, health, part-time status, migration, marital status, urbanicity, and (in some models) the presence of children.

Blinder (1973, p. 441) frames the problem somewhat differently:

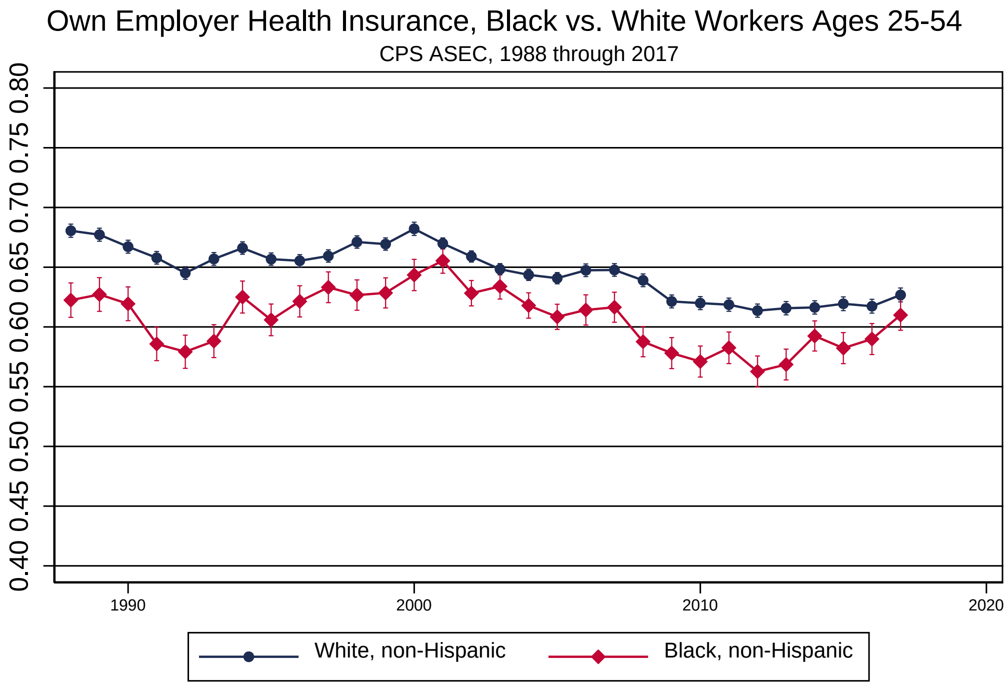
> In the intuitive model I have in mind, each individual is presented with endowments of human and non-human capitals and at some point in his life-cycle, jointly decides how far he wishes to pursue his formal education and to what occupational strata he aspires. Thus [education] and [occupation], the two chief determinants of the wage rate, are endogenous and simultaneously determined.

Blinder's solution is to use family background variables to address this endogeneity. Again, from Blinder (1973, p. 440):

> … the Michigan [Panel Study of Income Dynamics] data provide a rich set of variables pertaining to the individual's family background. These enable us to estimate a meaningful reduced-form equation which explains the wage rate only on the basis of characteristics which are truly exogenous to the individual (such as his father's education).

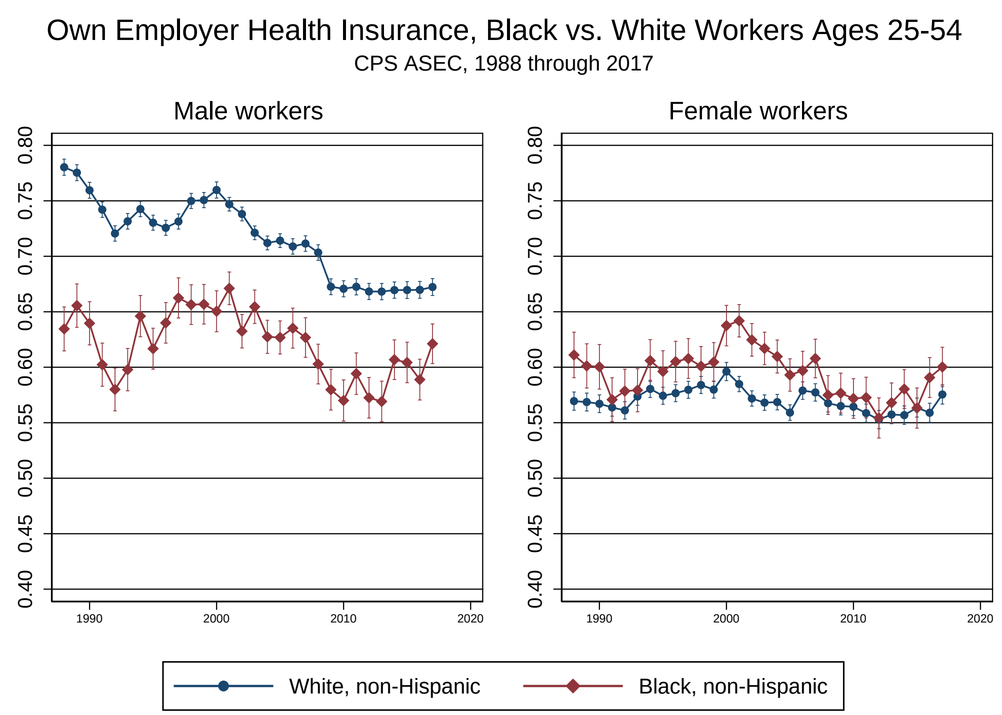
In addition to father's education, Blinder's list of exogenous variables includes (among others) age, region, and parents' income. The inclusion of parents' income and father's education in this list may strike the modern reader as odd, since it is inconceivable that parents' income and father's education would not have been affected by race and racism. Indeed, the 1967 data used by Blinder (1973) might have included some workers whose parents were born as enslaved persons. The choice of these variables arises in part due to how Blinder frames the problem. He was interested in separating discrimination from other factors, conditional on "the circumstances of [the worker's] birth" (p. 442) – implicitly starting the clock on racism at that point and ignoring what came before. This is consistent with the cultural and legal framework at the time, which focused on understanding whether employers were engaging in active discrimination based on race when faced with otherwise equally qualified applicants. At the same time, nearly 50 years later, it is clear that framing the problem in this way precludes any role for the long-lasting impacts of centuries of systemic racism in the United States and is therefore unsatisfactory if the goal is to understand the full impact of race on labor market outcomes.

Figure 1



Note: Vertical line segments show 95% confidence intervals.

Figure 2



Note: Vertical line segments show 95% confidence intervals.

Figure 3

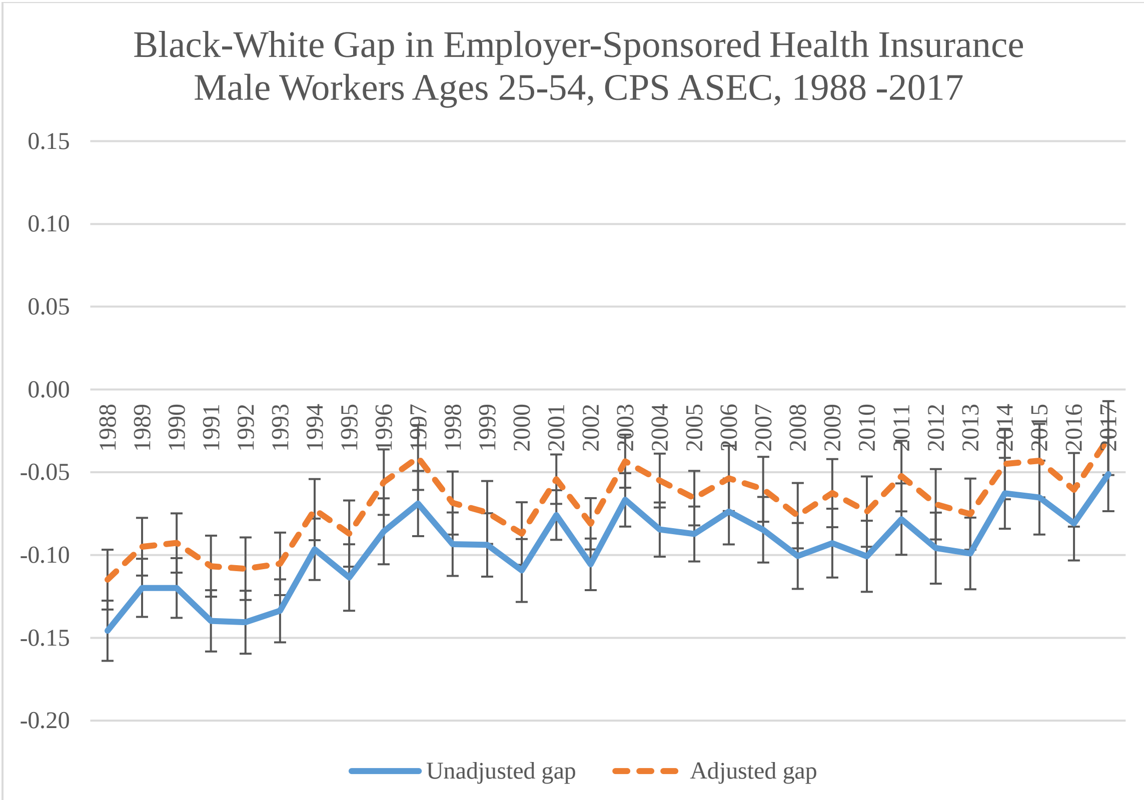
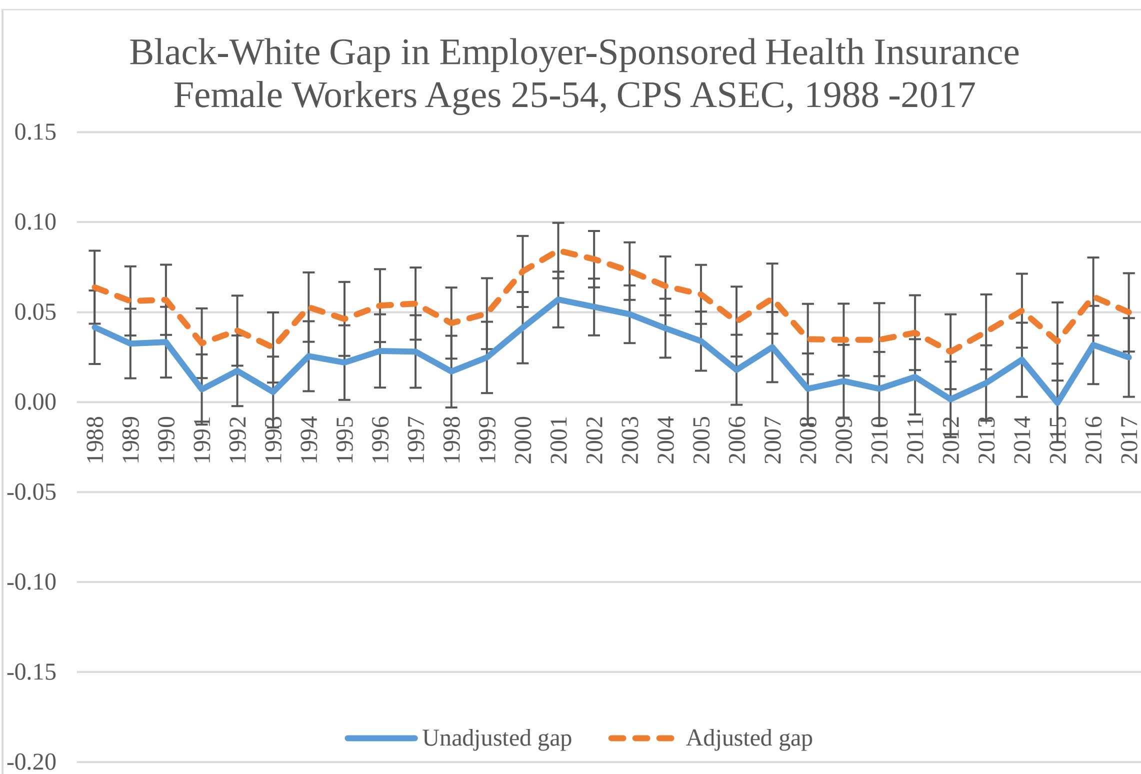


Figure 4



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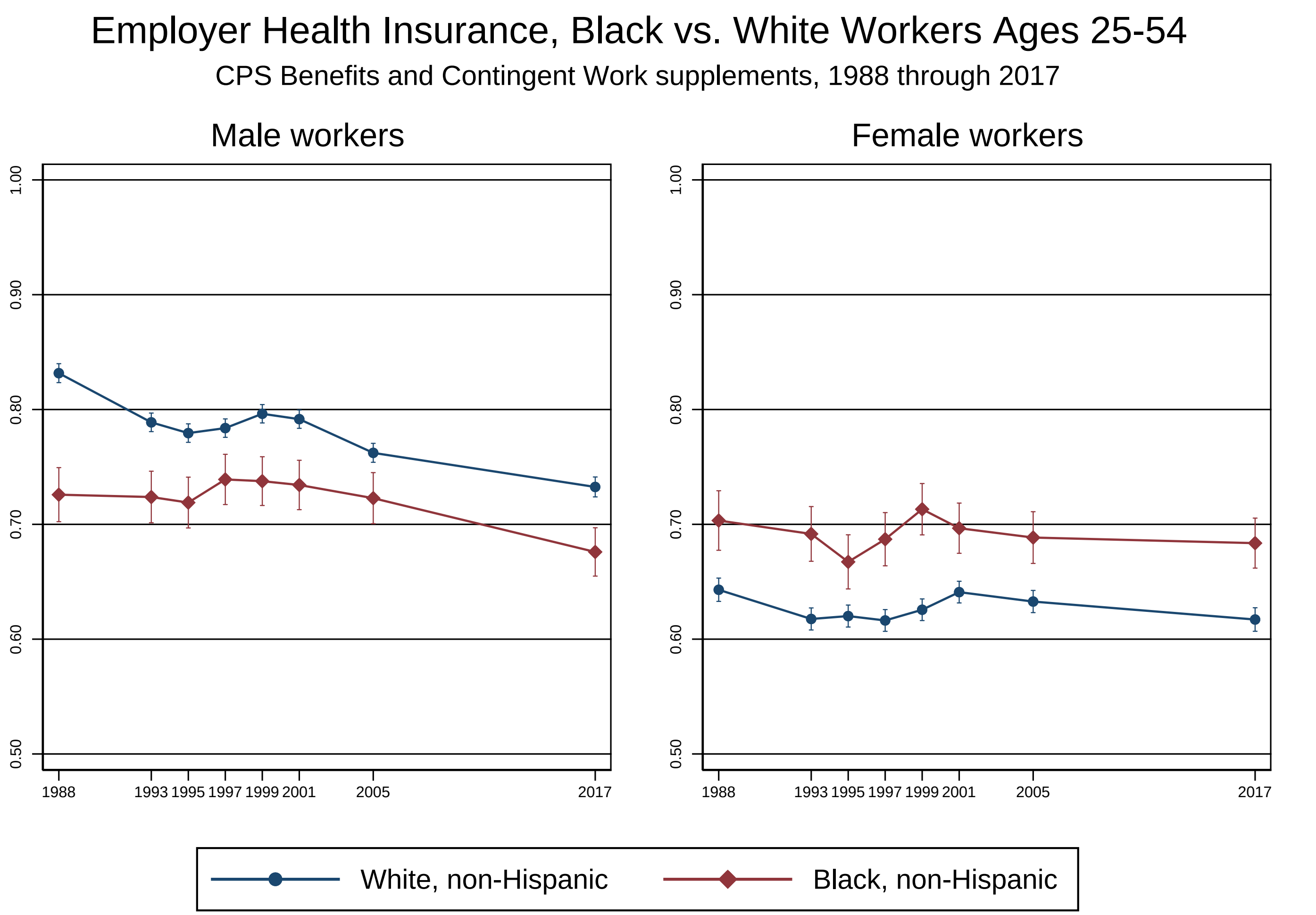

Note: Vertical line segments show 95% confidence intervals.

Figure 6

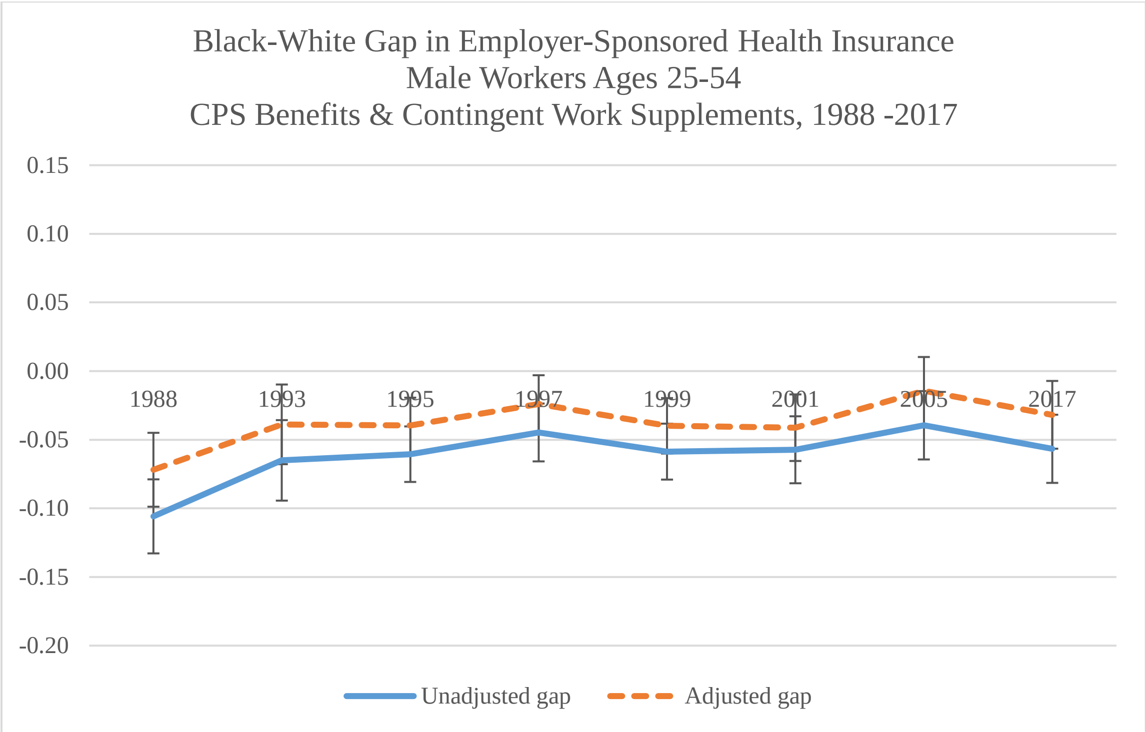

Figure 7

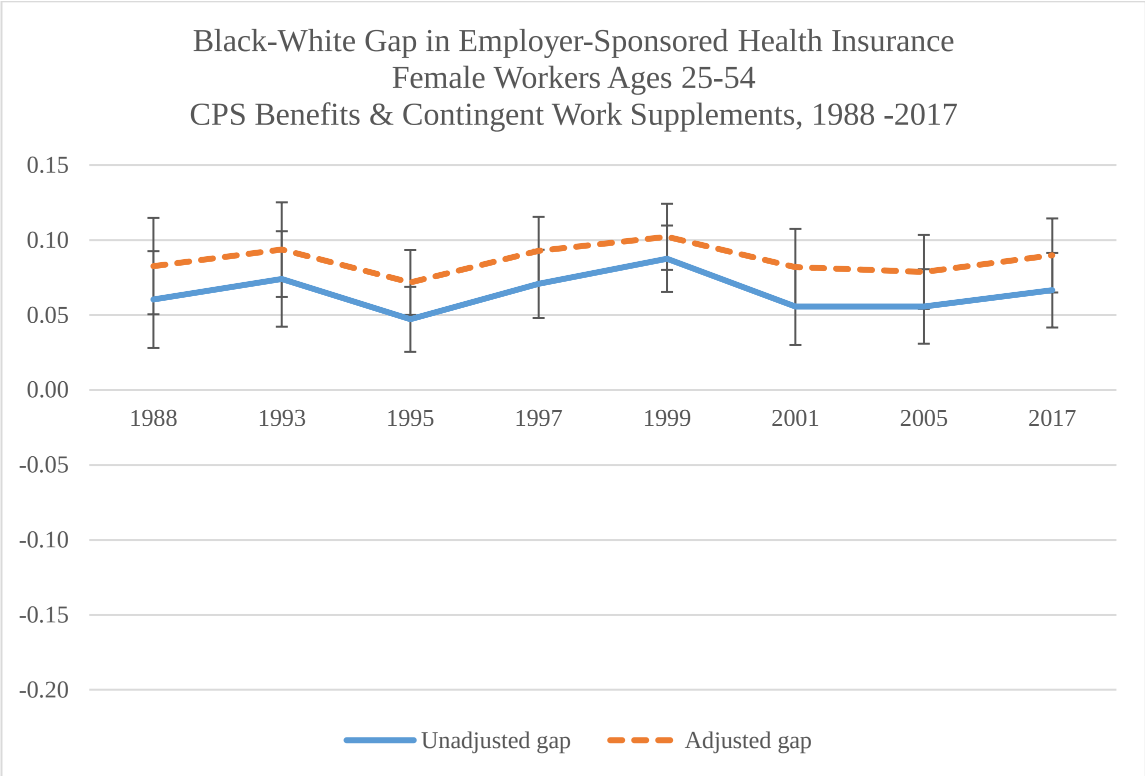

Figure 8

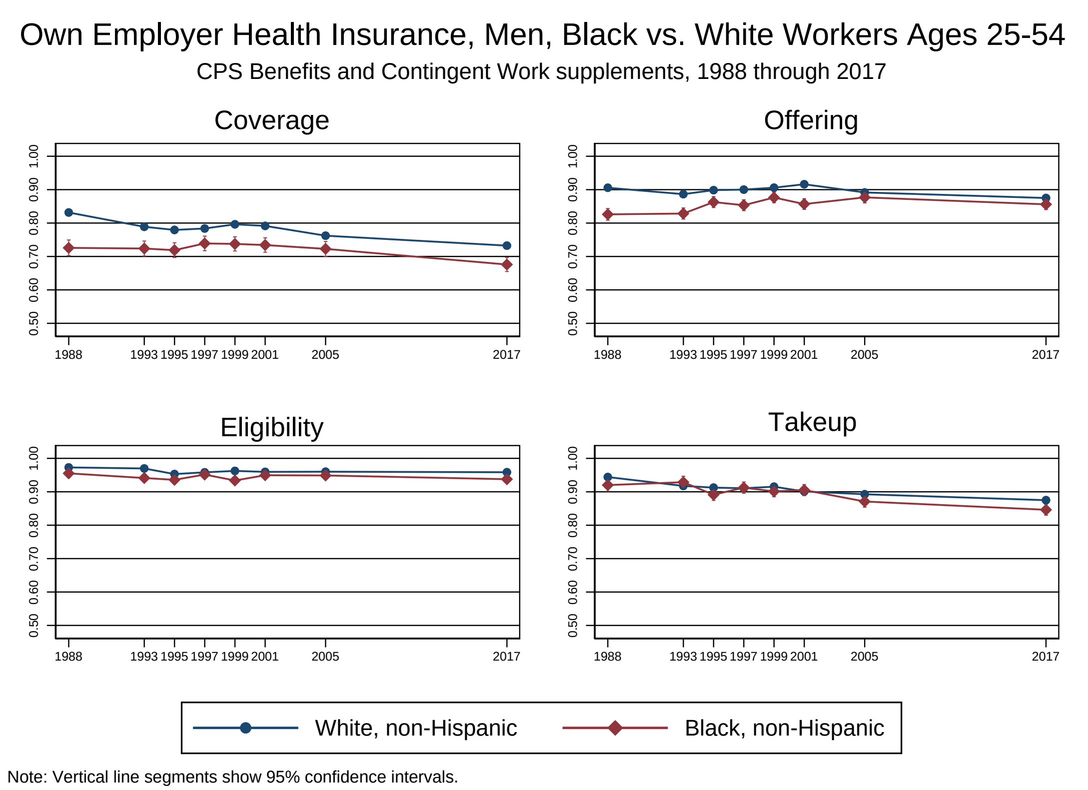

Figure 9

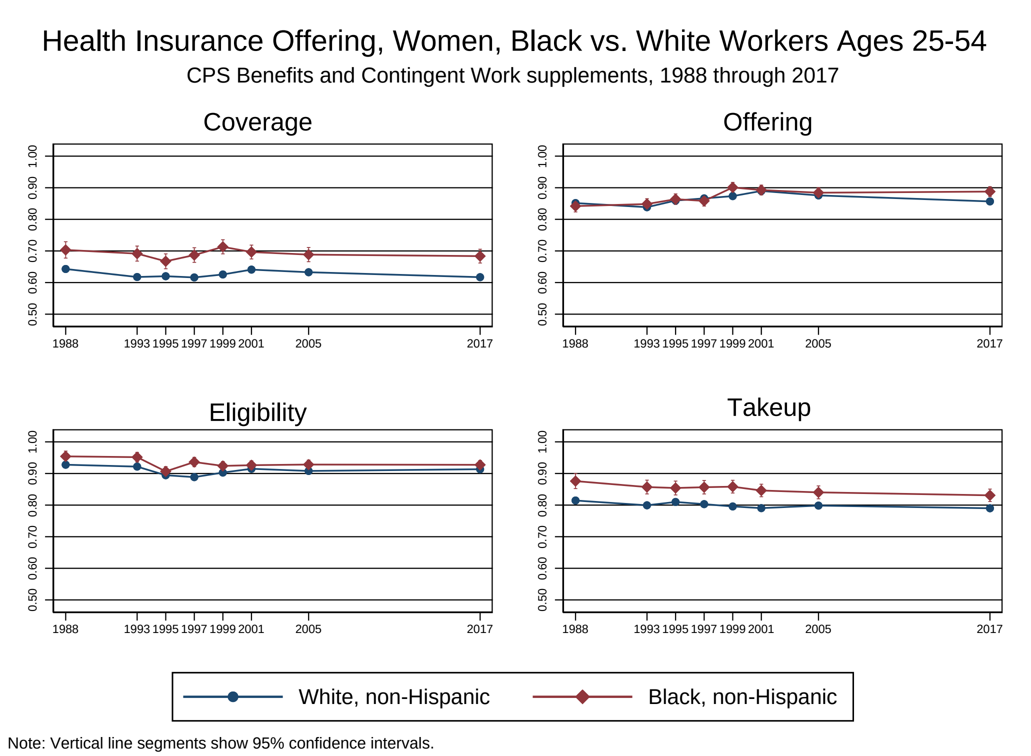

Figure 10

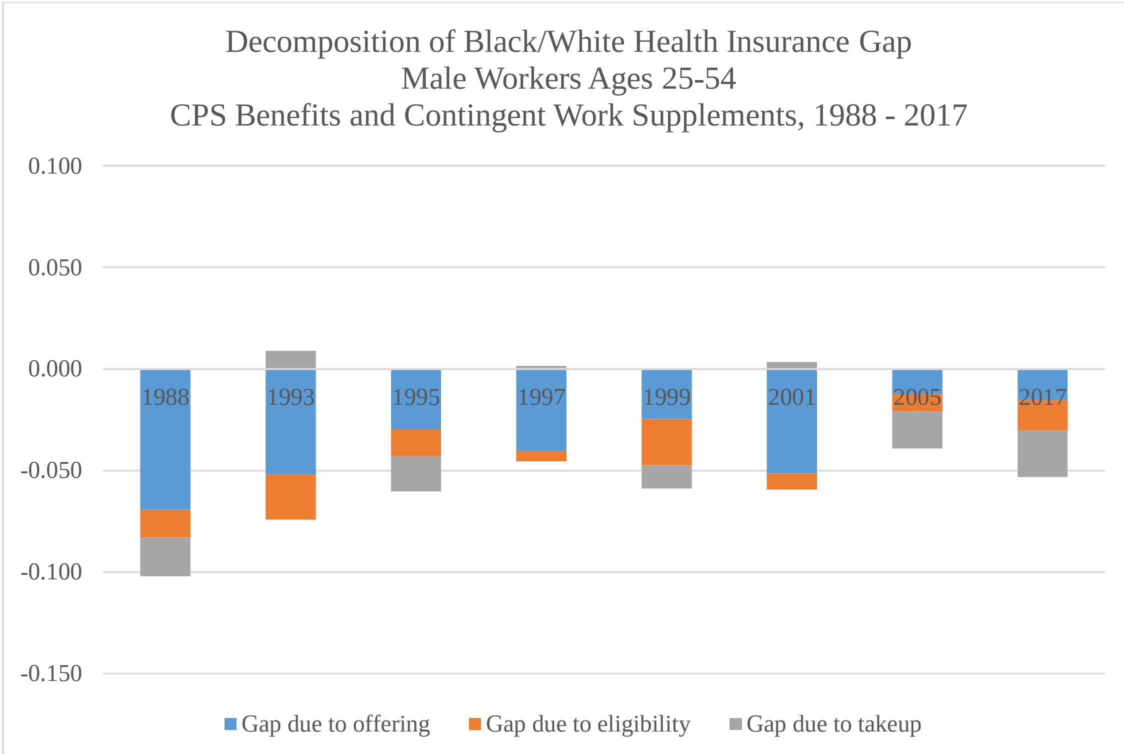

Figure 11

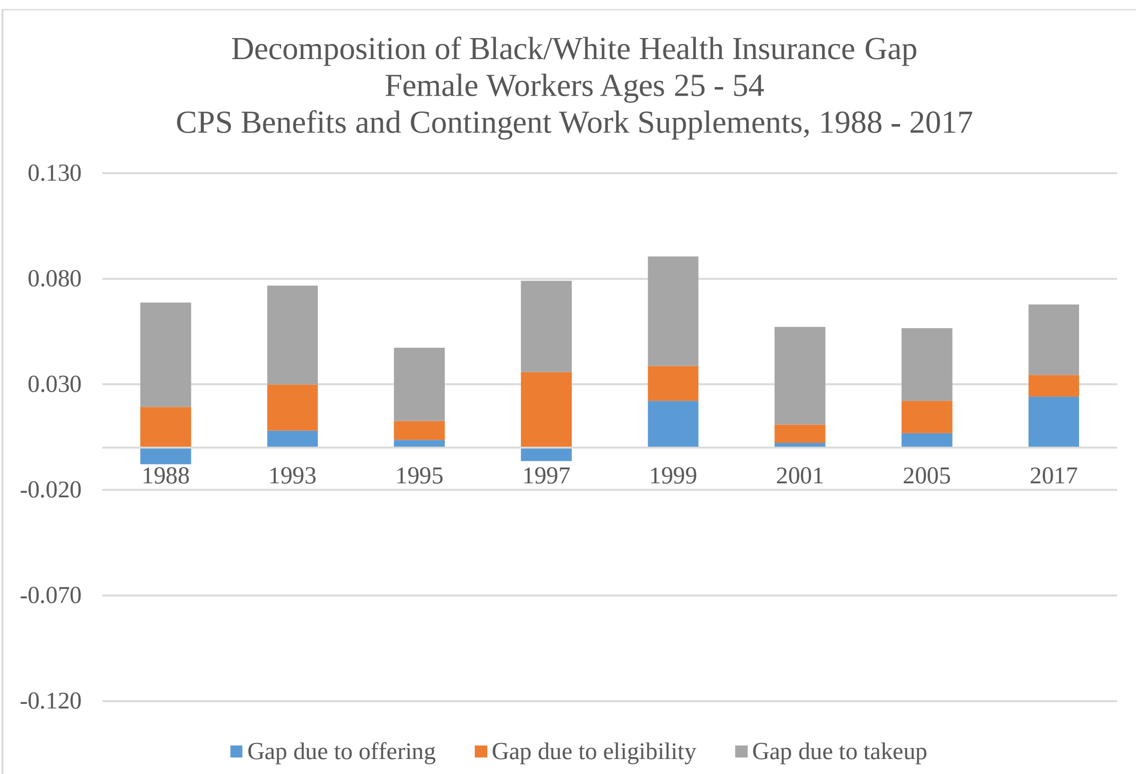

Table 1
Characteristics of Black and White Workers, by Race and Gender
CPS Annual Social and Economic Supplement, 1988 - 2017 (pooled)

	White men	Black men	White women	Black women
Own-employer health insurance	0.718	0.623	0.570	0.597
Any health insurance	0.883	0.782	0.900	0.820
Education < high school graduate	0.060	0.105	0.042	0.088
Education = high school graduate	0.310	0.386	0.290	0.331
Education = some college	0.270	0.297	0.301	0.337
Education ≥ college	0.359	0.211	0.367	0.244
Age	39.3	38. 5	39.4	38.6
Married	0.674	0.499	0.648	0.361
Public sector	0.137	0.177	0.198	0.240
Full-time, full-year workers	0.837	0.783	0.655	0.718
Part-time, full-year workers	0.024	0.035	0.122	0.073
Full-time, part-year workers	0.118	0.149	0.128	0.143
Part-time, part-year workers	0.021	0.034	0.095	0.066
Firm size 1 – 49	0.254	0.194	0.239	0.150
Firm size 50 – 99	0.129	0.111	0.115	0.099
Firm size ≥ 100	0.602	0.683	0.630	0.740
Sample n	613,623	75,497	569,980	95,512

Note: Firm size data are missing for approximately 16% of the sample in 1988 through 1990.

Table 2
Characteristics of Black and White Workers, by Race and Gender
CPS Benefits/Contingent Work Supplements
1988, 1993, 1995, 1997, 1999, 2001, 2005, and 2017 (pooled)

	White men	Black men	White women	Black women
Own-employer health insurance	0.785	0.723	0.628	0.692
Employer offers insurance	0.898	0.855	0.864	0.874
Eligible for insurance (if offered)	0.962	0.944	0.909	0.931
Takes up insurance (if eligible)	0.909	0.896	0.800	0.851
Any health insurance	0.901	0.827	0.891	0.841
Education < high school graduate	0.055	0.097	0.040	0.080
Education = high school graduate	0.321	0.401	0.315	0.345
Education = some college	0.271	0.301	0.296	0.337
Education ≥ college	0.353	0.201	0.349	0.237
Age	38.8	38.0	39.1	38.3
Married	0.719	0.601	0.680	0.485
Public sector	0.155	0.206	0.213	0.267
Job tenure (mean, in years)	8.236	7.015	6.775	6.934
Job tenure < one year	0.137	0.156	0.160	0.171
Sample n	82,574	7,954	78,473	10,423

Note: data on job tenure are missing for about 3.3% of respondents.

Table 3
Decomposition of Black-White Health Own-Employer Insurance Gaps among Workers
into Gaps Due to Offering, Eligibility, and Takeup
CPS Benefits/Contingent Work Supplements
1988, 1993, 1995, 1997, 1999, 2001, 2005, and 2017 (pooled)

	White Mean	Black Mean	Difference (Black-White)	Std. error of difference	Decom-position
Panel A: Male and female workers (pooled)					
Has own-employer coverage	0.710	0.707	-0.004	0.003	
Offered coverage	0.882	0.865	-0.016	0.002	
Eligible if offered	0.937	0.937	0.000	0.002	
Takeup conditional on eligibility	0.860	0.872	0.012	0.003	
Gap in coverage					-0.004
Gap due to offering					-0.013
Gap due to eligibility					0.000
Gap due to takeup					0.010
Panel B: Male workers					
Has own-employer coverage	0.785	0.723	-0.062	0.004	
Offered coverage	0.898	0.855	-0.043	0.003	
Eligible if offered	0.962	0.944	-0.018	0.002	
Takeup conditional on eligibility	0.909	0.896	-0.013	0.003	
Gap in coverage					-0.062
Gap due to offering					-0.036
Gap due to eligibility					-0.014
Gap due to takeup					-0.011
Panel C: Female workers					
Has own-employer coverage	0.628	0.692	0.064	0.005	
Offered coverage	0.864	0.874	0.010	0.003	
Eligible if offered	0.909	0.931	0.022	0.003	
Takeup conditional on eligibility	0.800	0.851	0.051	0.004	
Gap in coverage					0.064
Gap due to offering					0.008
Gap due to eligibility					0.017
Gap due to takeup					0.042

Note: The sample includes wage and salary workers between the ages of 25 and 54 who are Black non-Hispanic or White non-Hispanic. Decomposition methods are described in the text.

Table 4
Decomposition of Black-White Health Own-Employer Insurance Gaps among Workers into Gaps Due to Offering, Eligibility, and Takeup by Year, 1988 - 2017
CPS Benefits/Contingent Work Supplements

	1988	1993	1995	1997	1999	2001	2005	2017
ALL WORKERS								
Gap in coverage	-0.031	-0.001	-0.013	0.008	0.009	-0.005	0.003	0.003
Gap due to offering	-0.040	-0.021	-0.013	-0.023	-0.001	-0.022	-0.002	0.005
Gap due to eligibility	0.001	-0.001	-0.004	0.014	-0.005	-0.002	0.002	-0.004
Gap due to takeup	0.007	0.020	0.004	0.017	0.014	0.018	0.004	0.002
MALE WORKERS								
Gap in coverage	-0.106	-0.065	-0.062	-0.045	-0.060	-0.056	-0.040	-0.055
Gap due to offering	-0.069	-0.052	-0.030	-0.040	-0.025	-0.051	-0.012	-0.015
Gap due to eligibility	-0.014	-0.022	-0.013	-0.005	-0.023	-0.008	-0.009	-0.015
Gap due to takeup	-0.019	0.009	-0.017	0.001	-0.012	0.003	-0.018	-0.023
FEMALE WORKERS								
Gap in coverage	0.060	0.074	0.046	0.070	0.087	0.056	0.055	0.066
Gap due to offering	-0.008	0.008	0.004	-0.006	0.022	0.002	0.007	0.024
Gap due to eligibility	0.019	0.022	0.009	0.036	0.016	0.009	0.015	0.010
Gap due to takeup	0.049	0.047	0.035	0.043	0.052	0.046	0.034	0.033

Note: The sample includes wage and salary workers between the ages of 25 and 54 who are Black non-Hispanic or White non-Hispanic. Decomposition methods are described in the text.

Table A1
CPS ASEC Supplements
Sample size and distribution by race and gender
Wage and salary workers ages 25 – 54

	N	White men	Black men	White women	Black women
1988	39,132	0.467	0.046	0.431	0.057
1989	42,992	0.463	0.049	0.427	0.060
1990	42,616	0.461	0.048	0.433	0.057
1991	43,703	0.466	0.047	0.429	0.058
1992	43,056	0.467	0.047	0.429	0.057
1993	41,801	0.466	0.045	0.431	0.058
1994	41,602	0.465	0.048	0.429	0.058
1995	36,735	0.468	0.046	0.427	0.059
1996	37,178	0.466	0.048	0.427	0.059
1997	36,511	0.463	0.048	0.429	0.061
1998	36,488	0.461	0.049	0.427	0.063
1999	36,499	0.459	0.051	0.425	0.065
2000	34,651	0.459	0.052	0.423	0.065
2001	60,845	0.449	0.057	0.420	0.073
2002	58,737	0.451	0.055	0.422	0.072
2003	56,504	0.449	0.055	0.424	0.072
2004	55,037	0.448	0.056	0.425	0.071
2005	53,383	0.450	0.055	0.420	0.074
2006	52,506	0.446	0.059	0.417	0.077
2007	52,018	0.445	0.059	0.418	0.077
2008	51,976	0.447	0.059	0.417	0.077
2009	50,609	0.444	0.062	0.416	0.078
2010	48,090	0.446	0.060	0.417	0.077
2011	46,540	0.449	0.060	0.414	0.076
2012	46,133	0.448	0.062	0.412	0.077
2013	44,930	0.446	0.064	0.412	0.078
2014	44,125	0.440	0.067	0.411	0.082
2015	40,639	0.439	0.069	0.405	0.087
2016	40,674	0.442	0.068	0.406	0.083
2017	38,902	0.440	0.069	0.405	0.086

Table A2
CPS Benefits and Contingent Work Supplements
Sample size and distribution by race and gender
Wage and salary workers ages 25 – 54

	N	White men	Black men	White women	Black women
1988	15,218	0.478	0.043	0.427	0.053
1993	14,863	0.466	0.039	0.443	0.052
1995	31,682	0.460	0.043	0.439	0.058
1997	28,203	0.462	0.042	0.440	0.056
1999	28,029	0.457	0.045	0.437	0.060
2001	20,133	0.460	0.046	0.432	0.062
2005	21,524	0.452	0.042	0.449	0.056
2017	19,772	0.454	0.053	0.428	0.066

Table A3
Regression Results: Dependent variable = Own-employer health insurance
Black Non-Hispanic and White Non-Hispanic Wage and Salary Workers in CPS-ASEC, 1988 and 2017

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Women				Men			
	1988		2017		1988		2017	
Black	0.042***	0.064***	0.025*	0.050***	-0.146***	-0.115***	-0.051***	-0.029**
	(0.010)	(0.010)	(0.011)	(0.011)	(0.009)	(0.009)	(0.011)	(0.011)
Education:								
< High school		Omitted		Omitted		Omitted		Omitted
= High school		0.123***		0.086**		0.173***		0.167***
		(0.014)		(0.029)		(0.010)		(0.022)
Some college		0.173***		0.149***		0.197***		0.213***
		(0.015)		(0.029)		(0.011)		(0.022)
College +		0.275***		0.264***		0.263***		0.293***
		(0.014)		(0.028)		(0.011)		(0.022)
Constant	0.569***	0.401***	0.575***	0.382***	0.780***	0.589***	0.672***	0.443***
	(0.004)	(0.013)	(0.005)	(0.028)	(0.003)	(0.009)	(0.004)	(0.021)
Observations	18,932	18,932	12,611	12,611	19,864	19,864	13,120	13,120

Standard errors in parentheses
=* p<0.05 ** p<0.01 *** p<0.001