Racial and Ethnic Differences in Nonwage Compensation*

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Abstract

Previous research has found that, after controlling for test scores, measured black-white wage gaps are small but unemployment gaps remain large. This paper complements this previous research by examining the incidence of employer-provided benefits from the same premarket perspective. However, marriage rates differ substantially by race, and the possibility of health-insurance coverage through a spouse's employer therefore distorts how the distribution of benefits available in the market to an individual is expressed in the distribution of benefits received. Two imputation strategies are used to address this complication. The evidence suggests that benefit availability gaps are small.

One of the most striking and influential developments in the literature on racial inequality in labor markets is Neal and Johnson's (1996) finding that measured black-white wage gaps are small when controls are included for math and reading skills acquired before entry into the labor market. Neal and Johnson's result suggests that wage gaps between blacks and whites in the U.S. labor market can be traced largely to differences in the acquisition of human capital or other characteristics correlated with test scores that can be observed early in life. Further research has qualified, but not overturned this result (e.g., Carneiro, Heckman, and Masterov, 2005).

A study by Ritter and Taylor (2011) finds that the same is not true of lifetime experience of unemployment. Although black-white unemployment gaps narrow substantially after controlling for Armed Forces Qualification Test (AFQT) scores and other premarket factors, they remain very large—greater than 50 percent. Thus the empirical evidence points to a stylized picture of small black-white wage gaps and large unemployment gaps.

In both of these papers, white-Hispanic gaps are smaller. After controlling for for AFQT scores, male white-Hispanic wage gaps essentially disappear, but unemployment gaps remain moderately large. The latter is true for women as well.

The present study complements those mentioned above by examining the availability of employer-provided benefits from the same premarket perspective. Since benefits are a large component of compensation for some workers and contribute to inequality (Pierce, 2001), conclusions about the effects of race in the labor market generally, and about the link between race and compensation specifically, are incomplete unless the role of nonwage compensation is considered.

Neal and Johnson's results suggest that the labor market is not an important source of racial income inequality—that this inequality is largely driven by factors in place before labor market experience commences—but Ritter and Taylor's findings about unemployment undercut this interpretation. The omission of nonwage compensation from this line of research leaves open the possibility that the labor market tends to place comparable black and white workers into jobs with similar monetary compensation, while sorting black workers disproportionately into jobs that do not offer important fringe benefits. This sorting could reflect some aspect of employers' choices or differing worker preferences about the mix of wage and nonwage compensation. A stark version of the former possibility is embodied in some dual labor market models in which discrimination keeps minorities out of primary sector jobs (Bulow and Summers, 1986).

Nonwage compensation cost is positively correlated with earnings, but a number of factors break the link between earnings and nonwage compensation. First, most benefits are employer-specific rather than job-specific, and within employers benefits are not linked, or only partially linked, to earnings. Second, the employer's cost of offering specific benefits can vary among employers. Third, individuals vary in their preferences about the mix of wage and nonwage compensation, and employers vary the mix they offer as a means of managing their human resources.

Royalty and Abraham (2006) draw attention to a fourth factor that is especially pivotal for understanding differences among women: Many individuals can obtain health insurance through their spouses' employers, often at a marginal cost that is very low relative to obtaining health insurance in the private market. This option is valuable, and the theory of compensating differentials implies that the option induces some individuals to alter their choice of job, choosing employment offers with no health insurance but with, for example, higher monetary compensation or fewer hours. Because marriage rates differ dramatically by race, this phenomenon obscures possible differences in the underlying offer distributions.[1] To be concrete, black women are more likely than white women to report that their employers offer benefits, but I argue below that this is a consequence of the spousal coverage option—white women are much more likely than black women to be married and, therefore, far more likely to obtain health insurance coverage via their husband's employer, allowing them to choose jobs that are superior on other dimensions.

Existing research on employer-provided benefits and race is concentrated on health insurance, where the focus is mainly on coverage. Levy (2006), using Current Population Survey data for 1995–2001, finds that black-white gaps in the offering of health insurance are not statistically significant after controlling for various individual characteristics; Hispanic-white gaps are about 5 percentage points. Levy also addresses the question of health insurance take-up. Dushi and Honig (2005), using data from the Survey of Income and Program Participation, separate single workers, married workers in one-earner households, and married workers in dual-earner households, thus reaching more complex conclusions. Both of these papers found that nearly all of the black-white and Hispanic-white differences are due to differences in offering rates and very little due to differences in take-up.

The present research differs from these previous studies in three important ways. First, it takes a premarket perspective, which, simplifies a number of issues, as outlined by Neal and Johnson (1996). The premarket perspective on benefit availability essentially estimates the probability of being offered a benefit given (only) initial conditions (characteristics of the individual prior to labor market entry), rather than conditioning on individual characteristics at the time of the survey and/or employer characteristics. Second, the focus of this paper is on benefits as compensation, rather than on understanding differences in health insurance coverage rates (the central issue of Levy's and Dushi and Honig's papers). One consequence of the different focus is that I study a wider range of benefits. Third, I take account of how marriage affects inter-racial gaps in benefit availability.

The National Longitudinal Survey of Youth 1979 (NLSY79) used here collects information on whether particular benefits are available to employed respondents. I concentrate on the three most important: health, retirement, and vacation benefits. These data are described in section 1. Section 2 discusses further the consequences of the spousal coverage option, and section 3 proposes two simple imputation strategies to assess the importance of the spousal coverage option.

Section 4 presents results that point to two conclusions regarding the pattern of nonwage compensation. First, cognitive skills, as measured by AFQT score, are consistently an important predictor of the availability of health, retirement, and vacation benefits; benefits do not differ from earnings in this respect. Second, for men measured black-white and Hispanic-white gaps disappear after controlling for AFQT. Results are similar in some ways for women, but muddied by the prevalence of health coverage via spouse's employer. Application of the imputations helps to bound the size of this effect.

Section 5 applies the theory of compensating differentials to address the possibility of selection biases arising from differential censoring at the participation margin and from racial preference differences about the mix of wage and nonwage compensation.

1 Data

1.1 Sample

The data come from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79) and exclude the NLSY79's supplemental sample of poor non-black/non-Hispanic individuals. In 1994 and 2004, the years for which benefit availability is analyzed, respondents were 29–37 and 39–47 years old, respectively.

Whether a particular benefit is available to a given worker is determined by whether the employer offers the benefit at all and, if so, whether the employee is eligible. For most employers, eligibility of an employee is determined by the number of hours worked. However, the sorting of individuals into benefits-eligible positions is part of the question the premarket perspective attempts to answer, as it may be part of a discriminatory process. Therefore, *all* employed individuals, including part-timers, the self-employed, and unpaid family workers are included in the analysis.[2]

Individuals living outside the 50 U.S. states and District of Columbia at the relevant interview date were excluded. For a small number of workers, nearly all men, type of residence is recorded as "jail." Although there are a number of ways an incarcerated individual could correctly report working, these cases were dropped.

Following other papers that have studied the premarket determinants of racial differences in the labor market, I use the NLSY79 classification of respondents into black, Hispanic, and non-black/non-Hispanic groups. For brevity I refer to as "white," though it includes Asians and a small number of Native Americans.

1.2 Benefits data

The NLSY79 asks a battery of questions about the availability of various employer-provided fringe benefits.[3] I concentrate on three that are both common and financially important: health, retirement, and vacation benefits. (Retirement benefits exclude employer contributions for social security.) In 1994 and 2004, these three benefits accounted for 76 and 77 percent of nonwage compensation voluntarily provided to non-federal employees (Bureau of Labor Statistics, 1994 and 2004). There is no information in the NLSY79 on the quantity or value of health and retirement benefits, so the analysis concerns only availability of benefits. The other types of benefits tracked in the NLSY79 are fairly rare (e.g., child care), of modest financial value (life insurance), or are of significant value to only a small fraction of employees (child care).

The questions regarding health insurance and retirement are:

> Did your employer make available to you...
>
> > Medical, surgical, or hospital insurance that covers injuries or major illnesses off the job?
> >
> > A retirement plan other than social security?

These questions refer clearly to availability rather than receipt of these benefits; indeed the words "make available" are emphasized on the questionnaires in some years. Because employees frequently must share the cost of these benefits, take-up rates are potentially an important consideration, but I do not address them here.

The survey asks directly about vacation days: "How many days of paid vacation [are/were] you entitled to each year?"[4] A small number of respondents who received dual-purpose vacation/sick days (129 in 2004) were coded as having vacation.

All of the analysis is conducted for both 1994 and 2004, using information for the respondent's primary current job. Prior to 1994, the benefits questions were not asked of respondents who worked less than 20 hours per week at their main jobs. Given this screen, 1994 provides the closest comparability to Neal and Johnson's paper, which used 1990-91 wages.

Table 1 summarizes the availability of the three benefits under study for full-time workers in the NLSY79. Two features of the table are notable in the present context. First, the overall incidence of individual benefits shown in the top panel is quite high, especially considering that these figures include part-time and self-employed workers.

A more important aspect of Table 1 is the last line, which indicates that these three benefits tend to cluster—if one is available, it is likely that all three are available; for each year/sex combination between 71 and 81 percent of respondents were offered either none of these benefits or all of them. For brevity, therefore, only two binary dependent variables are used: availability of health insurance and availability of all three benefits.

1.3 Premarket covariates

The argument for using only covariates measured prior to labor market entry when addressing racial disparities is that variables measured later may be contaminated by the effects of race in the labor market (Neal and Johnson, 1996). Deciding whether a variable was measured premarket is trivial in some cases, but requires judgement in other cases.

The Armed Forces Qualification Test (AFQT), which is of central interest here, falls in the latter category. The AFQT score is extracted from a larger set of aptitude tests (the Armed Services Vocational Aptitude Battery), incorporating tests of mathematical knowledge, arithmetical reasoning, paragraph comprehension, and word knowledge. Because AFQT scores rise with age, the score is adjusted for age and standardized.

Neal and Johnson (1996) defined premarket variables conservatively and thus omitted individuals born before 1961 on the grounds that they were over 18 when the AFQT was administered to NLSY respondents in 1980. Neal and Johnson argued that once individuals enter the labor market, the AFQT score may be affected by labor market experience. Accordingly, they restricted their sample to individuals born after 1961. The point is correct in principle, but appears to make little difference in practice (other than for standard errors, of course), so I do not impose this restriction.[5]

Three other less important variables raise the same issue. First, the Rosenberg self-esteem scale "describes a degree of approval or disapproval toward oneself." (Center for Human Resources Research, 2001, p. 105). Data for this scale were also collected in 1980. It has 24 distinct values, and higher values correspond to greater self-esteem.

Second, the abbreviated Rotter internal-external locus of control scale measures "the extent to which individuals believe they have control over their lives through self-motivation or self determination (internal control) as opposed to the extent that the environment (i.e., chance, fate, luck) controls their lives (external control)." (Center for Human Resources Research, 2001, p. 104). The scale was administered in 1979. It has 13 distinct values, and higher values correspond to greater external focus.

Third, because arrest records can have significant labor market effects, I use information collected in 1980 about whether the respondent had ever been charged with illegal activity and whether he or she was ever convicted.

The remaining covariates unambiguously meet the premarket standard. These are mother's and father's education, as measured by highest grade completed, and the number of times the respondent was suspended or expelled from school. The latter variables are intended to capture attitudes toward school and human capital accumulation.

2 Compensating differentials and household decisions

One of the peculiar institutional characteristics of the health insurance system in the U.S. is that employer-provided health insurance will usually cover an employee's family and, moreover, premiums are usually subsidized for family members. This means that the value of an offer of health insurance coverage has far less value to certain married workers because they can be covered under their spouse's employer's plan. Because health insurance is so costly in the United States, this oddity can provide a strong incentive for married individuals to select jobs that substitute higher pay, more leisure, or other amenities for redundant health coverage.

Figure 1 shows a simple variation on the standard textbook model of compensating differentials induced by fringe benefits. An indifference curve for an individual with health insurance from her own employer is labeled *OO*, with point A being the optimal point on the offer curve. Point B corresponds to the same compensation package with health insurance removed.

Suppose now that this individual becomes eligible to be covered at no cost under her spouse's health insurance (identical to her own). The new circumstance changes her preferences in this space (the horizontal axis represents benefits paid by *her* employer). Point A on the offer curve is still available and results in the same utility level as before, but the marginal rate of substitution between wages and benefits is different: the marginal value of health insurance offered by her employer is lower because of her new health insurance option. The indifference curve with through point A, labeled *SS*, is therefore flatter. It passes to the left of point B (since the worker has health insurance and a higher wage not offset by a premium).

In the new circumstances the worker chooses a point like C, which corresponds to a different mix of benefits and wages. Of course the model abstracts from the reality that the menu of fringe benefits is not smooth or consistent from employer to employer. The key point, however, is that an individual who is married has the option to "sell" her health insurance offer by moving to a different point on her offer curve.[6] Some workers, such as the one described by figure 1, will exercise the option.

Since most employers do not have "cafeteria" plans, but instead offer fixed packages of benefits, a worker who chooses a job without health insurance benefits tends forgo other important benefits as well, for which spousal options are not available. On the other hand, part of the gain at point C could be be additional flexibility, the opportunity to work part time, etc.

The magnitude of the phenomenon just described is suggested by table 2, which shows the prevalence of coverage via spouse's employer for NLSY79 respondents who were employed in 1994. The patterns are similar for 2004 (but not reported). The "not offered but covered by spouse's employer" lines are the most interesting. These are people who might have taken advantage of the option of spousal coverage in order to take a job that is preferred on other dimensions. The second row indicates that 17 percent of white women were not offered health insurance at their own jobs, but obtained coverage via spouse's employer. That rate was far higher than the rates for black and Hispanic women. Among men, the corresponding rates are much lower, but were more than twice as high for whites as for blacks or Hispanics.

These differences can be traced partly to different marriage rates, but are proportionately much larger; the propensity for married people to follow this option differs by race. For example, white women in the NLSY79 were only twice as likely to be married in 1994 as black women, but 3.5 times more likely to appear in the "not offered but covered by spouse's employer" line of table 2.

If the probability of coverage via spouse's employer is related to race and that in turn changes the characteristics of an individual's job (i.e., the move from point A to point C), racial gaps in benefit availability *on the current job* will differ systematically from opportunities to obtain benefits through job search. Specifically, availability of benefits on the current job will be suppressed more for whites than for blacks and Hispanics.

3 Imputation strategies

The option of health coverage through a spouse's employer means that some individuals will be offered fewer benefits by their chosen employers than are available to them in the market. The ideal imputation would identify these people and assign them benefits. Unfortunately, though we can easily identify the people for whom this imputation *might* be needed, there is no obvious way to determine which of those with spousal coverage could actually find a job with benefits. This section proposes two imputations intended to help assess the impact of the spousal coverage option on measured racial differences.

The process of assortative mating leads to spouses with similar characteristics. This suggests that individuals with spousal coverage would generally be able to find jobs with benefits if they chose to do look for them. That line of reasoning leads to:

Assortative-Mating Imputation: Assign benefit availability to *all* employed workers with spousal coverage.

Imputation based on assortative mating was also used by Neal (2004), who argued that most non-working married women with a prosperous spouse could command a high wage if they chose to work and imputed high wages to them (in a median regression).[7]

Obviously, the true probability of being able to find a job with benefits conditional on one's spouse having benefits is less than one. However, Royalty and Abraham argue that unobservable characteristics are likely to be highly correlated with those of a respondent's spouse and this is an important advantage of this imputation.

Unfortunately, the validity of that argument likely varies by race because average characteristics of individuals with imputations differ by race. For example, average AFQT score of white women with spousal health coverage is about one standard deviation higher than that of black or Hispanic women. This fact would tend to cause the assortative mating imputation to understate disparities; the white women's higher AFQT scores mean they would have a higher probability of finding a job with health insurance if they chose to look for one.

A second, complementary approach is to bound the effect of spousal coverage on racial disparities. Recall that the reason for considering these imputations is that white workers are more likely to use the spousal coverage option. The impact of that difference cannot be larger than implied by the following imputation:

Bounding Imputation: Assign benefit availability to all white workers with spousal coverage.

Do not assign benefits to non-white workers with spousal coverage.

All of the reported results for women also include an imputation for non-working, welfare-dependent women with no spousal income, which closely parallels a similar imputation used by Neal (2004). This imputation changes results only trivially and is not discussed further.

4 Results

4.1 Men

Table 3 presents linear probability models without imputations for men employed in 1994.[8]

The first and third columns indicate that black men were significantly less likely than white men to be offered either health insurance alone or the three-benefit package. However, the regressions in columns 2 and 4 show that differences in cognitive skills and other premarket factors (primarily the former) account for the entire gap. Unadjusted benefit availability gaps between white Hispanic men are smaller, but the qualitative effect of including AFQT and other regressors is the same.

Table 4 shows the effect of the two imputation strategies on health insurance availability. As expected, since not many men have insurance through their spouses' employers, these imputations do not make a great deal of difference for men. The regressions using bounding imputation (columns 5 and 6) find that black and Hispanic men are slightly less likely to have health benefits after controlling for premarket factors. Recall, however, that this imputation is constructed to make things look as bad as possible for non-whites; the imputation says that a disparity larger than 3.7 percentage points is not being masked by failing to account adequately for spousal coverage.

The overall conclusion for men, then, is that nonwage compensation behaves in much the same way as monetary compensation—inter-racial gaps in benefit offerings can be fully accounted for by differences in premarket factors.

Regressions based on 2004 benefits data (the regressors remain unchanged) are reported in an appendix. The key results are similar for individuals employed in 2004. However, disparities (before adding controls) between white and non-white workers are generally smaller in 2004. An interesting difference is that spousal coverage appears to matter more: If no imputation is used, after adding controls black and Hispanic men appear to be better off than white men. This phenomenon is much more pronounced in the data on women, to which I now turn.

4.2 Women

Linear probability models for women using no imputations are shown in Table 5. These regressions display what at first seems a striking anomaly: black and Hispanic women are considerably *more* likely to be offered benefits at their current jobs than white women, and the gap grows dramatically when controls for cognitive skills and other premarket factors are added. However, as argued earlier, although these regressions correctly estimate the actual coverage gaps among working women, because of the possibility of coverage through a spouse's employer, the measured gaps likely differ significantly from disparities in the offer distribution. As suggested by table 2, this is a more important issue for women than for men.

Before considering the imputations it is useful to note that in two important respects the women's results are consistent with the men's. First, AFQT score matters, just as it does for wages and unemployment. Second, adding the premarket covariates to the regressions moves the black-white and Hispanic-white gaps in the positive direction by approximately the same amount as for men.

The effects of the imputations are reported in table 6. Compared to table 4, the imputations have a dramatic effect on the measured black-white differential, reducing it to near zero with either imputation. Thus coverage through a spouse's employer plays an important role in causing the anomaly mentioned above. Most important, the results using the bounding imputation (columns 5 and 6), show that the spousal coverage option does not mask a significantly negative coefficient on black. A significantly negative impact of being Hispanic cannot be ruled out for women in 1994, though it can be in 2004 (table A-4).[9] Further 2004 results are presented in the appendix.

The evidence from the imputations speaks most directly to health insurance benefits, but it seems reasonable to extend it to all three important benefits for two reasons. First, the regressions that include vacation and retirement benefits are very similar to the health insurance regressions without imputations. Second, given the strong tendency for employers to offer benefit packages that include all three benefits, the dynamics of health insurance coverage are likely to spill over to other benefits. In other words, if an worker takes advantage of her husband's health insurance coverage by choosing a no-health-insurance job from her offer curve, she is likely to sacrifice retirement and vacation benefits as well.

Table 7 takes another approach by restricting the sample to only never-married individuals. The estimated gap between single white and black women is extremely close to zero when the full set of regressors is included, but the white-Hispanic gap is puzzlingly large and positive.

NJ pointed out that general equilibrium models such as Lundberg and Startz (1983) predict that minorities will underinvest in skills if discrimination lowers the return to those skills. NJ addressed this concern by interacting black with AFQT. They concluded that there was no evidence that returns to AFQT differed between blacks and whites. Following that lead I interacted AFQT with the dummies for black and Hispanic. The dependent variable was availability of health insurance including the bounding imputation, which is conservative with respect to the conclusion of no difference.

Wald tests of the null that all coefficients involving black (or Hispanic) do not alter the overall conclusion. The terms involving black are insignificant and the magnitudes of effects near the mean of AFQT are small. For male Hispanics, the coefficients are jointly significant, but the magnitude of the effect remains small.[10]

5 Endogeneity and Compensating Differentials

The evidence presented in section 4 supports the view that the offer distributions for benefits do not differ significantly among whites, blacks, and Hispanics. However, in estimating racial benefit gaps there are two sources of potential endogeneity that were not addressed. First, it is possible that preferences about the mix of wage and nonwage compensation differ by race. For example, black workers might put a higher value on benefits than white workers. This would lead black workers to select into jobs with benefits, partially concealing any racial differential in benefit availability.

Second, different participation rates of white and black (Hispanic) workers imply that their respective probabilities of receiving benefits might be differentially biased by censoring at the participation margin. Previous research has found that in wage regressions any bias induced by differential participation is not large enough to dramatically change the interpretation of Neal and Johnson's results (Neal and Johnson, 1996; Johnson, Kitamura and Neal, 2000). However, there is no parallel evidence regarding benefits, and methods used in the referenced papers, such as imputing zero wages to nonparticipants, cannot be used. There is no obvious way to address these issues directly in estimation, but the theory of compensating differentials sheds light on both questions.

First consider the possibility of different preferences about the mix of wage and nonwage compensation. In figure 2 black workers favor a mix of compensation tipped toward benefits compared to white workers facing the same offer curve. Neal and Johnson's result was that wages are approximately the same for both groups.[11]

The conventional interpretation of that result is that black and white workers face the same offer curve, but in figure 2 it only says that the tangencies lie close the same horizontal line. The present paper finds that benefits are approximately equal, so that the tangencies should lie close to the same vertical line.[12] That combination of results is geometrically impossible unless the expansion paths are close to each other, i.e., that preferences do not differ much. This line of reasoning does not depend on which group prefers more benefits.

The argument can be stated in terms of compensating differentials. The diagram is drawn so that black workers put a higher value on benefits than white workers. That induces them to sort into jobs with benefits. This sorting would offset the effects of discrimination, potentially producing results similar to those I report here. But that scenario is inconsistent with Neal and Johnson's results: the higher benefits received by black workers imply a negative compensating differential.

What if black workers place less value on benefits than white workers (switching the expansion paths in Figure 2)? In this case discrimination would reinforce rather than offset the effect of preferences. Black workers would sort into jobs with fewer benefits, and they ought to see a positive compensating differential. That scenario is also inconsistent with the results reported here, that estimated benefit gaps are near zero.

Turn now to the issue of differential selection at the participation margin. Figure 3 is how figure 2 looks if whites and blacks prefer the same mix of wages and benefits. In the figure, preferences over wages and benefits are homothetic so that the expansion path is linear, given a set of hypothetical offer curves, but the argument below does not depend on this. Looked at in this way, individuals do not select a reservation wage or reservation benefit level, per se. Instead wages and benefits are determined jointly.[13] The reservation utility level that must be reached by a job offer ($U = U^*$) is achieved at a point on the expansion path. The individual chooses to work if that point is below the offer curve they actually face (point A in figure 3).

In this framework the standard selection concern is that preferences might differ in such a way that one racial group tends to choose reservation points much farther out on their expansion path, so that participation rates differ significantly in the two groups. Discrimination, on the other hand, amounts to moving the offer curve toward the origin, also producing movement along the expansion path. Obviously the two effects are easily confounded.

The theory of compensating differentials emphasizes that individuals choose wages and benefits jointly; faced with the offer curve shown in figure 3, the individual will apply for jobs with a mix of wages and benefits on or near the expansion path. Thus, whatever bias is induced by censoring at the participation margin applies to both wages and benefits. Therefore, evidence that the bias in wage regressions is small implies that the bias would be small in benefits regressions as well. (This is why the imputations used in section 4 do not assign benefits to non-working individuals covered by their spouses' insurance—the evidence about the size of selection bias in wage equations would not apply with a different sample selection process.)

6 Conclusion

This paper studies nonwage compensation from the premarket perspective. It finds that nonwage benefits behave similarly to wages in important respects. First, cognitive skills, as measured by AFQT scores, are an important and highly significant predictor of benefit availability. Second, controlling for cognitive skills changes estimated black-white and Hispanic-white benefit availability gaps in the direction advantageous to the minority group. A qualification to these conclusions is that only the availability of benefits, not their value, was studied.

After making imputations that account for the possibility of obtaining health benefits via a spouse's employer, there is little evidence that white-nonwhite gaps in health insurance availability are are very different from zero in either practical or statistical terms.

Combining this finding with those of Neal and Johnson (1996) and Ritter and Taylor (2011) suggests a broad-brush description of the U.S. labor market: appropriately measured compensation gaps between races are small, while unemployment gaps are large.

Notes

1. This point is closely related to an argument made by Neal (2004) that inter-racial demographic differences distort the black-white wage gap for women. It also hinges on the fact that for married individuals the source of health insurance coverage is endogenous to household decision making (Royalty and Abraham, 2006).

2. This is also why 1994 was the first year used: it is the first year in which all working individuals are asked about benefits. Before 1994 only those working 20 or more hours were asked. The 20-hour screen would create a selection bias that is correlated with race, especially for women as a higher fraction of white women work part time. See section 2.

3. Individuals who report working less than 10 hours per week are first asked whether they are offered any benefits. If they respond in the affirmative, they are asked the same series of questions as those working more hours.

4. One of the unexpected problems of asking this question directly is that it raises the frequency of "don't know" responses—nearly 2 percent in 2004 vs. 0.2 percent for health insurance. This is presumably because some respondents do not know their exact entitlement, while knowing they have some vacation. Similarly, the "don't know" rate for the NLSY question about sick leave, which is handled in the same way, is 6 percent for 2004.

5. Neal (2005) and Ritter and Taylor (2011) make the same choice.

6. This option also has implications for the wage distribution, which are implicitly addressed by Neal (2004).

7. Although my imputation follows the same general logic, the imputations here are for employed individuals, rather than non-employed ones. Section 5 discusses selection at the participation margin.

8. The analysis was performed using R version 2.15 (R Development Core Team, 2012).

9. The values of the regressors do not change between Tables 6 and A-4. The samples differ somewhat, but if the regressions are restricted to only those respondents who answered the health benefits question in both years, the comparison between the years is similar.

10. For 2004, the Wald test indicates that terms involving Hispanic are significant for women, while the coefficient is insignificant in table A-4.

11. The gap for black men is about 1.5 log points larger if 1994 wages are used in place of 1990-91 wages and is about the same fraction of the raw gap (before controling for AFQT). The coefficient on Hispanic is very close to zero for 1994.

12. Strictly speaking, I find that benefit *availability* is approximately equal. It is conceivable that black workers, while having access to the same benefits as white workers, get lower value benefits, though regressing number of vacation days on the same variables produces similar results (not reported).

13. This is not equivalent to choosing a reservation compensation level since utility is not necessarily monotonic in the cost of compensation to the employer, precisely because workers have preferences over the mix of wage and nonwage compensation.

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Table 1: Benefit Availability (unweighted percent)

	1994		2004	
	Men	Women	Men	Women
Overall incidence				
Vacation	73.7	68.8	73.3	70.1
Health	71.7	65.8	73.5	70.8
Retirement	56.2	54.7	64.6	65.7
Incidence of combinations				
None	19.3	25.3	21.2	22.3
Vacation	7.7	7.4	4.3	5.1
Health	3.2	1.8	1.6	1.1
Retirement	0.6	0.6	0.4	1.0
Vacation+Health	13.5	11.0	8.3	6.7
Vacation+Retirement	0.8	1.4	0.9	1.9
Health+Retirement	3.2	3.7	3.6	5.5
Vacation+Health+Retirement	51.7	48.8	59.8	56.3

Table 2: Health Insurance Availability and Coverage, 1994 (Percent)

	Hispanic	Black	White
Women			
Not offered	35.0	30.5	36.0
but covered by spouse's employer	11.3	4.8	17.0
Offered	65.0	69.5	64.0
but covered by spouse's employer	13.4	11.2	16.8
Men			
Not offered	30.6	32.1	25.5
but covered by spouse's employer	1.9	2.0	4.4
Offered	69.4	67.9	74.5
but covered by spouse's employer	4.8	4.3	7.1

Notes: Percent of all employed individuals (not restricted to regression samples).

Table 3: Linear Probability Models of Benefit Availability, Men, 1994

	Health Insurance		Health Insurance, Retirement, and Vacation	
Intercept	0.754****	0.634****	0.553****	0.492****
	(0.010)	(0.091)	(0.011)	(0.103)
Black	−0.079****	0.008	−0.073****	0.034
	(0.018)	(0.021)	(0.020)	(0.022)
Hispanic	−0.053**	0.018	−0.046**	0.038
	(0.021)	(0.023)	(0.023)	(0.025)
Age		0.010***		0.010**
		(0.003)		(0.004)
AFQT		0.077****		0.092****
		(0.010)		(0.011)
$AFQT^2$		0.007		0.003
		(0.007)		(0.008)
Mother's education		−0.000		0.001
		(0.002)		(0.002)
Father's education		0.002		0.001
		(0.002)		(0.002)
Self-esteem scale		0.006***		0.008****
		(0.002)		(0.002)
Rotter scale		0.001		−0.002
		(0.004)		(0.004)
Suspensions		−0.002		−0.005***
		(0.003)		(0.002)
Expulsions		−0.005		−0.016
		(0.015)		(0.012)
Charged with illegal activity		−0.138****		−0.129****
		(0.030)		(0.030)
Convicted of illegal activity		−0.002		0.014
		(0.038)		(0.040)
$\bar{R}^2$	0.006	0.057	0.004	0.062
N	3458	3458	3458	3458

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1
Heteroskedasticity corrected standard errors in parentheses.

Table 4: Health Insurance Availability with Imputations, Men, 1994

	No imputation		Assortative mating		Bounding	
Covariates included?	No	Yes	No	Yes	No	Yes
Intercept	0.754****	0.634****	0.795****	0.690****	0.795****	0.723****
	(0.010)	(0.091)	(0.009)	(0.085)	(0.009)	(0.087)
Black	−0.079****	0.008	−0.096****	−0.009	−0.118****	−0.037*
	(0.018)	(0.021)	(0.017)	(0.020)	(0.018)	(0.020)
Hispanic	−0.053**	0.018	−0.068****	0.002	−0.090****	−0.024
	(0.021)	(0.023)	(0.020)	(0.022)	(0.020)	(0.022)
$\bar{R}^2$	0.006	0.057	0.010	0.065	0.015	0.068
N	3458	3458	3522	3522	3522	3522

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

Heteroskedasticity corrected standard errors in parentheses.

Table 5: Linear Probability Models of Benefit Availability, Women, 1994

	Health Insurance		Health Insurance, Retirement, and Vacation	
Intercept	0.641****	0.711****	0.483****	0.542****
	(0.012)	(0.100)	(0.012)	(0.105)
Black	0.063****	0.141****	0.059***	0.153****
	(0.019)	(0.021)	(0.020)	(0.023)
Hispanic	0.014	0.080***	−0.003	0.083***
	(0.023)	(0.025)	(0.024)	(0.026)
Age		−0.001		−0.002
		(0.004)		(0.004)
AFQT		0.066****		0.086****
		(0.011)		(0.011)
$AFQT^2$		−0.021***		−0.023***
		(0.008)		(0.008)
Mother's education		−0.000		0.003
		(0.002)		(0.002)
Father's education		0.001		−0.001
		(0.002)		(0.002)
Self-esteem scale		0.007***		0.007***
		(0.002)		(0.002)
Rotter scale		0.005		0.006
		(0.004)		(0.004)
Suspensions		−0.020***		−0.020***
		(0.007)		(0.007)
Expulsions		−0.007		0.001
		(0.016)		(0.015)
Charged with illegal activity		−0.062		−0.057
		(0.059)		(0.058)
Convicted of illegal activity		0.026		0.026
		(0.092)		(0.091)
$\bar{R}^2$	0.003	0.032	0.002	0.040
N	3265	3265	3265	3265

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1
Heteroskedasticity corrected standard errors in parentheses.

Table 6: Health Insurance Availability with Imputations, Women, 1994

	No imputation		Assortative mating		Bounding	
Covariates included?	No	Yes	No	Yes	No	Yes
Intercept	0.641****	0.711****	0.810****	0.697****	0.810****	0.716****
	(0.012)	(0.100)	(0.009)	(0.087)	(0.009)	(0.090)
Black	0.063****	0.141****	−0.084****	0.030	−0.128****	−0.015
	(0.019)	(0.021)	(0.017)	(0.019)	(0.017)	(0.020)
Hispanic	0.014	0.080***	−0.047**	0.051**	−0.158****	−0.062***
	(0.023)	(0.025)	(0.019)	(0.021)	(0.021)	(0.023)
$\bar{R}^2$	0.003	0.032	0.007	0.069	0.026	0.080
N	3265	3265	3401	3401	3401	3401

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

Heteroskedasticity corrected standard errors in parentheses. As described in section 3, columns 3–6 include imputations for welfare-dependent, non-working women.

Table 7: Health Insurance Availability, Single Individuals, 1994

	Men		Women	
Intercept	0.653****	0.706****	0.741****	0.333
	(0.024)	(0.184)	(0.028)	(0.209)
Black	−0.070**	0.047	−0.109***	−0.001
	(0.034)	(0.041)	(0.038)	(0.044)
Hispanic	−0.008	0.046	0.032	0.125**
	(0.046)	(0.049)	(0.049)	(0.052)
Age		0.001		0.024***
		(0.008)		(0.008)
AFQT		0.099****		0.093****
		(0.019)		(0.022)
$AFQT^2$		0.001		0.001
		(0.015)		(0.016)
Mother's education		−0.003		−0.000
		(0.004)		(0.004)
Father's education		0.002		0.000
		(0.003)		(0.003)
Self-esteem scale		0.002		0.008*
		(0.004)		(0.004)
Rotter scale		−0.006		0.009
		(0.008)		(0.009)
Suspensions		0.002		−0.023
		(0.003)		(0.024)
Expulsions		−0.014		−0.020
		(0.040)		(0.079)
Charged with illegal activity		−0.100*		−0.231*
		(0.056)		(0.132)
Convicted of illegal activity		−0.043		−0.040
		(0.081)		(0.213)
$\bar{R}^2$	0.003	0.039	0.014	0.090
N	988	988	712	712

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

OLS estimates with heteroskedasticity corrected standard errors in parentheses.

Appendix

Table A-1: Linear Probability Models of Benefit Availability, Men, 2004

	Health Insurance		Health Insurance, Retirement, and Vacation	
Intercept	0.757****	0.587****	0.631****	0.420****
	(0.011)	(0.100)	(0.012)	(0.112)
Black	−0.057***	0.015	−0.059***	0.046*
	(0.019)	(0.022)	(0.021)	(0.024)
Hispanic	−0.001	0.056**	−0.023	0.056**
	(0.022)	(0.024)	(0.025)	(0.027)
Age		0.011***		0.012***
		(0.004)		(0.004)
AFQT		0.066****		0.090****
		(0.011)		(0.012)
$AFQT^2$		0.008		−0.003
		(0.008)		(0.009)
Mother's education		−0.000		0.001
		(0.002)		(0.003)
Father's education		0.000		0.000
		(0.002)		(0.002)
Self-esteem scale		0.004**		0.006**
		(0.002)		(0.002)
Rotter scale		0.004		0.008*
		(0.004)		(0.005)
Suspensions		−0.004		−0.008***
		(0.004)		(0.003)
Expulsions		−0.007		−0.014
		(0.017)		(0.018)
Charged with illegal activity		−0.137****		−0.158****
		(0.034)		(0.034)
Convicted of illegal activity		−0.003		0.065
		(0.043)		(0.044)
$\bar{R}^2$	0.003	0.041	0.002	0.054
N	2879	2879	2879	2879

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1
Heteroskedasticity corrected standard errors in parentheses.

Table A-2: Health Insurance Availability with Imputations, Men, 2004

	No imputation		Assortative mating		Bounding	
Covariates included?	No	Yes	No	Yes	No	Yes
Intercept	0.757****	0.587****	0.805****	0.588****	0.805****	0.618****
	(0.011)	(0.100)	(0.010)	(0.094)	(0.010)	(0.095)
Black	−0.057***	0.015	−0.079****	0.002	−0.104****	−0.030
	(0.019)	(0.022)	(0.018)	(0.021)	(0.018)	(0.022)
Hispanic	−0.001	0.056**	−0.038*	0.026	−0.053**	0.005
	(0.022)	(0.024)	(0.021)	(0.023)	(0.021)	(0.023)
$\bar{R}^2$	0.003	0.041	0.006	0.051	0.011	0.051
N	2879	2879	3015	3015	3015	3015

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

Heteroskedasticity corrected standard errors in parentheses.

Table A-3: Linear Probability Models of Benefit Availability, Women, 2004

	Health Insurance		Health Insurance, Retirement, and Vacation	
Intercept	0.702****	0.691****	0.562****	0.561****
	(0.012)	(0.103)	(0.013)	(0.111)
Black	0.034*	0.122****	0.058***	0.162****
	(0.019)	(0.022)	(0.021)	(0.024)
Hispanic	−0.010	0.070***	0.001	0.091****
	(0.023)	(0.025)	(0.025)	(0.027)
Age		0.004		−0.001
		(0.004)		(0.004)
AFQT		0.068****		0.090****
		(0.011)		(0.012)
$AFQT^2$		−0.040****		−0.042****
		(0.008)		(0.008)
Mother's education		0.001		0.000
		(0.003)		(0.003)
Father's education		0.001		0.001
		(0.002)		(0.002)
Self-esteem scale		0.002		−0.000
		(0.002)		(0.002)
Rotter scale		−0.005		0.000
		(0.004)		(0.005)
Suspensions		−0.007*		−0.003
		(0.004)		(0.004)
Expulsions		−0.107***		−0.119****
		(0.038)		(0.030)
Charged with illegal activity		−0.094		−0.149**
		(0.058)		(0.058)
Convicted of illegal activity		−0.044		0.003
		(0.100)		(0.097)
$\bar{R}^2$	0.001	0.040	0.002	0.043
N	2929	2929	2929	2929

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1
Heteroskedasticity corrected standard errors in parentheses.

Table A-4: Health Insurance Availability with Imputations, Women, 2004

	No imputation		Assortative mating		Bounding	
Covariates included?	No	Yes	No	Yes	No	Yes
Intercept	0.702****	0.691****	0.820****	0.697****	0.820****	0.732****
	(0.012)	(0.103)	(0.010)	(0.087)	(0.010)	(0.091)
Black	0.034*	0.122****	−0.048***	0.073****	−0.086****	0.021
	(0.019)	(0.022)	(0.016)	(0.018)	(0.017)	(0.020)
Hispanic	−0.010	0.070***	−0.053***	0.057***	−0.127****	−0.033
	(0.023)	(0.025)	(0.020)	(0.021)	(0.021)	(0.023)
$\bar{R}^2$	0.001	0.040	0.003	0.075	0.015	0.069
N	2929	2929	3123	3123	3123	3123

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

Heteroskedasticity corrected standard errors in parentheses. As described in section 3, columns 3–6 include imputations for welfare-dependent, non-working women.

Table A-5: Health Insurance Availability, Single Individuals, 2004

	Men		Women	
Intercept	0.691****	0.363	0.804****	0.302
	(0.034)	(0.266)	(0.038)	(0.268)
Black	−0.146***	−0.050	−0.129***	0.002
	(0.046)	(0.057)	(0.048)	(0.055)
Hispanic	−0.060	−0.013	−0.013	0.103
	(0.060)	(0.066)	(0.065)	(0.071)
Age		0.015		0.021**
		(0.010)		(0.010)
AFQT		0.083***		0.063**
		(0.027)		(0.027)
$AFQT^2$		0.033		−0.052**
		(0.020)		(0.022)
Mother's education		−0.001		−0.001
		(0.006)		(0.006)
Father's education		−0.001		0.007*
		(0.004)		(0.004)
Self-esteem scale		0.001		−0.006
		(0.006)		(0.006)
Rotter scale		0.005		−0.004
		(0.011)		(0.011)
Suspensions		−0.001		−0.015
		(0.013)		(0.018)
Expulsions		−0.020		−0.127
		(0.044)		(0.131)
Charged with illegal activity		−0.139*		−0.273
		(0.083)		(0.171)
Convicted of illegal activity		−0.021		−0.078
		(0.113)		(0.354)
$\bar{R}^2$	0.015	0.033	0.014	0.090
N	532	532	432	432

Significance: **** = 0.001, *** = 0.01, ** = 0.05, * = 0.1

OLS estimates with heteroskedasticity corrected standard errors in parentheses.