SOCIAL CLASS, RENT DESTRUCTION, AND THE EARNINGS OF BLACK AND WHITE MEN, 1982–2000

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ABSTRACT

Adopting a consistent social class schema for March CPS data, we analyze trends from 1982 to 2000 in social class and black-white differences in the earnings of men. After presenting the well-known recent growth in earnings inequality within an EGP social class framework, we then show that the steady decline in the black-white gap among full-time, full-year workers is concentrated among semi-skilled and unskilled workers (classes VI and VIIa) and is particularly pronounced among supervisors of manual workers (a large proportion of class V). Three explanations account for this pattern. First, in the context of growing wage inequality across the entire labor market, white men at the bottom of the class distribution have been unable to maintain their wage advantages over their black counterparts to the same degree that white men at the top of the class distribution have. Second, rates of labor market non-participation among low-skilled black men increased in the 1980s and 1990s (though not as much as they increased in the 1960s and 1970s), and this selection out of the labor force has accentuated the decline in the relative earnings gaps observed for classes V, VI, and VIIa. Third, the tight labor market of the 1990s has disproportionately helped low-skilled black men. The relative weighting of these three explanations cannot

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be determined definitively with any available data, but the first explanation deserves at least as much attention as the latter two, has not been sufficiently developed in the literature on the black-white gap, and is consistent with the heretofore unevaluated rent destruction explanation for increases in inequality of Sørensen (2000). Taken together, the findings demonstrate the continued value of traditional modes of social class analysis for the study of inequality in post-industrial society.

INTRODUCTION

Examining social class differences in the earnings of black and white men from 1982 through 2000, we will show that the significance of race for labor market inequality among men has changed in the economic transformation of the 1980s and 1990s. In the course of analysis, we will argue that there has been a relative decline in the importance of race at the bottom of the earnings distribution of men and that this change can be explained in large part by general increases in earnings inequality. In developing this argument, we engage three related strands of literature – explanations for growth in earnings inequality in the U.S. labor market since the early 1980s (see Card & DiNardo, 2002; Katz & Autor, 1999), empirical examinations of the relative position of black men over the same time period (e.g. Grodsky & Pager, 2001; Wilson, 1987, 1996), and recent theoretical elaborations of the structural foundations of social class (Goldthorpe, 2000; Sørensen, 1996, 2000).[1]

PAST FINDINGS ON CHANGES IN INEQUALITY AND BLACK-WHITE DIFFERENCES IN EARNINGS

In the transformation of the U.S. economy in the 1980s and 1990s, labor market inequality increased dramatically across years of completed education and even more comprehensively within almost all measured categories of workers (see Katz & Autor, 1999). As a result, earnings inequality in the U.S. labor market is now more pronounced than perhaps at any time since World War II. Despite the theoretical importance and public recognition of these trends, relatively little sociological literature has investigated the causal mechanisms responsible for generating them (see Morris & Western, 1999). Even a descriptive agenda has not been pursued seriously, as there is little or no published literature on the growth in earnings gaps between social classes.[2]

Within the relevant economics literature, no strong evidence in support of a single explanatory narrative has emerged. Where there was some initial agreement, surrounding the explanation of skill-biased technological change (see Berman et al., 1994, 1998), Morris and Western (1999, p. 635) surmise that "something other than science" fortified the initial enthusiasm.[3] And, thus, in addition to technological change, many labor economists are willing to accept that an unknown mixture of demographic shifts in the early 1980s, globalization, deindustrialization, and associated declines in union membership and negotiating power also seem responsible for a substantial portion of the increase (see the SDI framework of Katz & Autor, 1999; see also Bernstein & Mishel, 2001; Card & DiNardo, 2002).

Alongside these more general changes in inequality, the black-white gap in the earnings of men has changed over the past several decades as well. Here again no explanation has achieved dominance, and the extant sociological literature is rather thin. Certainly sociologists have studied trends in black-white differences in occupational attainment (e.g. Hout, 1984) and in labor force participation (e.g. Jargowsky, 1997).[4] Again, however, economists have dominated the study of trends in black-white earnings differentials (see Bound & Freeman, 1992; Bound & Holzer, 1993; Card & Lemieux, 1994, 1996; Juhn et al., 1991). Altonji and Blank (1999) summarize the economics literature, highlighting several common empirical findings. Although the black-white earnings gap decreased dramatically between 1960 and 1980, it has since remained relatively stable. And for the last two decades, two countervailing trends must be separated. Black men continued to narrow some of the wage gap by acquiring more education, and perhaps also by securing higher occupational attainments above and beyond that which can be attributed to their increased educational attainments. But because the earnings of low-skilled workers have not kept pace with high-skilled workers over the same period and because black men still on average have lower educational attainments than white men, the gains won by increased educational attainment have been wiped out almost entirely in the aggregate.

Although this basic two-part narrative seems to have been widely accepted in the mid-1990s, considerable ferment emerged in the late 1990s over fine points that do and do not support it. The work collected in Cherry and Rodgers (2000), especially Freeman and Rodgers (2000), argues that the tight labor market of the 1990s disproportionately benefitted black men relative to their white counterparts. Not only did unemployment fall relatively more for blacks throughout the 1990s, but the wages of blacks appear to have increased relative to whites, particularly at the bottom of the distribution of educational attainment. Others, however, have countered these claims, arguing that race differences in patterns of incarceration (see Western & Pettit, 2000) and slight increases in relative non-employment among blacks (even in the context of declining unemployment) may have created illusory gains (see also Juhn, 2001). This counter-argument represents a re-birth of arguments that as much as 50% of the wage gains observed in the 1960s and 1970s for blacks represent artefactual increases attributable to gradually increasing labor market non-participation (see Brown, 1984; Heckman et al., 2000; Jaynes, 1990; Juhn, 1992; Smith & Welch, 1989; Welch, 1990).

WHAT ROLE FOR CLASS ANALYSIS?

To the extent that sociologists are more predisposed than economists to regard labor market positions as more fundamental than individuals' attributes, a *prima facie* sociological depiction of changes in inequality would focus on the linkage between earnings levels and alternative occupational positions, the latter being conceptualized broadly as social class position. And yet, as we noted earlier, a straightforward and sufficiently complete class-based descriptive analysis of recent trends in earnings inequality cannot be found in the literature. Should such an agenda be pursued?

Debates over the optimal form of class analysis for future research have proliferated in recent years (see Grusky & Weeden, 2001; Portes, 2000; Sørensen, 2000; along with their associated comments and replies). These debates are crucial for refining the conceptual foundations of research on social inequality, but they should not be allowed to crowd out empirical research with the tools at our disposal. New empirical work gives these debates more material with which to grapple, extending the terrain of contestation beyond the classic conjectures of the 1960s through 1980s on patterns of social fluidity and styles of life.

In this article, we aim primarily to move the empirical literature forward and hence will offer a straightforward social class decomposition of trends in earnings inequality, focusing narrowly on the earnings of black and white men (see Note 1). And, we will use only one of the social class schemas that is available, the one variously known as the EGP schema (after Erikson et al., 1979), Goldthorpe's class schema (after Goldthorpe, 1987), or the CASMIN coding (after Goldthorpe & Müller, 1982). Although many class mappings exist in the sociological literature, in the past two decades the EGP schema has become the most prominent, primarily because it has been effectively deployed in a wide variety of substantive contexts, such as in studies of social mobility (e.g. Erikson & Goldthorpe, 1992; Hout, 1989) and voting (e.g. Heath et al., 1985; Manza & Brooks, 1999). The EGP schema has also received a recent theoretical justification (see Goldthorpe, 2000) grounded on a broad set of literature from both economics and sociology, and has been introduced into the economics literature (see Erikson & Goldthorpe, 2002). Table 1 presents descriptions of the EGP social classes. Our analysis is based on our own coding of the 1980 and 1990 census occupation codes and is available from the authors by request.

Table 1. Descriptions for the EGP Social Class Schema.

Social Class	Description
I	Higher-grade professionals, administrators, and officials; managers in large industrial establishments; large proprietors
II	Lower-grade professionals, administrators, and officials; higher-grade technicians; managers in small industrial establishments; supervisors of non-manual employees
IIIa	Routine non-manual employees, higher-grade (administration and commerce)
IIIb	Routine non-manual employees, lower-grade (sales and service)
IVa and IVb	Small proprietors, artisans, and other self-employed workers
IVc	Farmers and small-holders and other self-employed workers in primary production
V	Supervisors of manual workers; lower grade technicians
VI	Skilled manual workers
VIIa	Semi- and unskilled manual workers (not in agriculture)
VIIb	Agricultural and other workers in primary production, not self-employed

Note: A detailed account of our coding of the class schema is provided in a supplementary appendix.

For our analysis, EGP classes serve as attractive vantage points from which to track the evolution of earnings differentials in the U.S. labor market. We do not regard the fine contours of labor market outcomes as the residue of overt class struggle, and hence we do not maintain that earnings are distributed according to the EGP class structure. Rather, the EGP class schema merely furnishes for us a useful grouping of structural positions in the labor market, clustered along specific (but not exhaustive) dimensions of employment relations.[5] We leave arguments for the general superiority of the EGP class schema, vis-à-vis its alternatives, to its creators (see Erikson & Goldthorpe, 2002; Goldthorpe, 2000), but we will offer some justification for its adoption in this application.

As a predictive tool, the EGP schema performs well in accounting for changes in inequality.[6] Indeed, as we will show later, had it been deployed in the late 1980s and early 1990s when economists were documenting movement in the college wage premium, sociologists may have discovered the growth in inequality.

And yet, predictive power, although attractive, is surely an insufficient justification on its own for deploying the EGP schema. We therefore offer four additional reasons for adopting the EGP schema for an analysis of changes in inequality in sociology:

(1) Why should sociology allow labor economists to define the explananda of interest for sociologists? The increase in the college wage premium is surely worthy of investigation, but narrowly focusing on the divergence between the mean earnings of classes I and V in the EGP framework is likewise interesting, especially when examining cross-cutting patterns of ascriptive inequality that have traditionally been framed with the language of social class.

(2) Although gross measures of inequality (e.g. Gini coefficients, Theil indices) will continue to be the dominant metric for cross-national comparisons of earnings distributions, EGP classes represent a more promising basis for first-order decompositions of any such cross-national differences. There appears to be considerably less variability across industrialized countries in the division of labor than in educational institutions. And, even more narrowly within the U.S., there is a great deal of unexamined heterogeneity, which arises from different types of educational institutions, within typical "years completed" measures of educational attainment. While EGP classes can be decomposed to the unit-occupation-level, with data such as the CPS no similar decomposition can be performed for different levels of educational attainment. Thus, while one can offer careful within-class adjustments for compositional change (as we will offer in Tables 3 and 4), no similarly direct adjustment can be offered for groups delineated by years of completed education.

(3) A class-based approach allows for a more explicit connection of studies of earnings inequality to traditional modes of analysis in social stratification research, primarily in social mobility studies (e.g. Erikson & Goldthorpe, 1992). Certainly, at some point in the future, we will want to answer the questions: Have these changes in earnings inequality affected regimes of social mobility? And if so, have the changes in mobility patterns been more dramatic in the U.S. where the increase in earnings inequality has been the greatest, perhaps thus lending support to the currently out-of-favor contention of American exceptionalism (see Erikson & Goldthorpe, 1985)?

(4) Narrative accounts of structural change are more naturally constructed using results generated by a class schema, especially when collective action and institutional constraints are important elements of the plot. Class analysis provides actors, albeit of an aggregated sort, that can be used to construct compelling accounts. How such narratives are best developed has become a contentious issue in the recent literature. For detailed "realist" narratives, Grusky and Sørensen (1998) and Grusky and Weeden (2001) have argued that aggregate actors considerably more narrowly defined than those suggested by the EGP schema are required. And yet, for sufficiently parsimonious and compelling narratives, Portes (2000) has argued that some sort of nominalist big class schema must be adopted, although not necessarily the EGP schema which we surmise he would regard as too rigidly bound by the occupational categories invented for national censuses and surveys.

For our application, we are satisfied that the EGP schema allows us to effectively and elegantly convey the descriptive findings of our analysis and hence serves justifications 1 through 3 outlined here. However, as will become clear in our discussion section, where we invoke the rent-based social class framework of Sørensen (1996, 2000), the EGP schema must also be supplemented conceptually in this particular application in order to account for within-class race differences. We do not regard the need for such supplementation as a fundamental limitation of the EGP class schema. Rather, to us, it demonstrates the utility of the schema, as it can accommodate many such types of supplementation, and hence allows for refinements in pursuit of a compelling explanatory narrative.

PLAN FOR EMPIRICAL ANALYSIS

Against this background of literature, we pursue answers to a series of descriptive questions:

(1) How has the black-white gap in the earnings of men evolved over the last several decades?
(2) How have EGP class differentials in the earnings of black and white men evolved since 1982?

In response, we show that a slight decline in the black-white gap has occurred in spite of generalized growth in earnings inequality across social classes. Attempting to reconcile these findings, we then ask:

(3) Is the decline in the black-white gap entirely attributable to changes in class entry rates?
(4) If not, is the black-white gap in earnings among men declining at the same modest rate within each class? Or, alternatively, is there interpretable variation in the decline across social class?

Although straightforward, these questions require careful modeling, especially when the limitations of the Current Population Surveys are recognized and the potential impact of self-selection out of the labor force is assessed. Nonetheless, the effort is justified, as the findings suggest important implications for the evolution of inequality in the late-twentieth century U.S. labor market. In the concluding section, we discuss these implications, drawing upon the rent destruction explanation for increases in inequality promoted by Sørensen (1996, 2000).

DATA AND VARIABLES

We analyze labor force data drawn from the 1964 through 2001 March Current Population Surveys. For the majority of our analysis, however, we restrict attention to the 1983 through 2001 March CPS data, and hence the analysis of prior-year earnings data from 1982 through 2000. We chose this more limited time frame to allow for a consistent and detailed coding of social class using 1980 and 1990 3-digit census occupation codes. Nonetheless, this nineteen-year time period is the focus of the literature on recent increases in earnings inequality and also nicely brackets the two most recent economic expansions (1983–1989 and 1992–2000, interrupted by the 1990–1991 downturn). In some of our analysis, we will explicitly compare the last five years of these two most recent economic expansions in an attempt to remove business cycle effects from our over-time comparisons. Presumably, at some point in the future, the 1990–1991 recession should be compared to the recession of the early 2000s, but we do not attempt such a comparison here (in part because it is unclear whether the latter has yet been completed). For all of our models, we analyze weekly earnings, converted to constant 2000 dollars using the Personal Consumption Expenditures Deflator (PCED). For the CPS, earnings values are top-coded to preserve the anonymity of respondents, and the top codes change periodically over the years. To create a uniform dataset, we imposed the lowest top-code for any year on all years, after having converted each year's values to 2000 dollars. We then multiplied the common topcode by a multiplier calculated using a method inspired by Pareto, which varied from 1.50 to 1.76 over the years.[7]

Our detailed coding of the EGP class schema (from the 1980 and 1990 census occupations codes and the employer size variable) is available in an extensive supplementary appendix that is available from the authors by request. For the regression models reported in Tables 3 and 4, we used standard codings of CPS variables for education, marital status, urbanicity, and region. Details of these coding decisions, along with those used to analyze more general patterns of labor force participation, are also available in the supplementary appendix.

RESULTS

The Evolution of Earnings Differentials Among Men in the 1980s and 1990s

We begin our empirical analysis by linking our findings to the established literature on the black-white gap in the earnings of men. Figure 1 presents five-year moving averages from 1963 to 2000 of the difference in log weekly earnings between

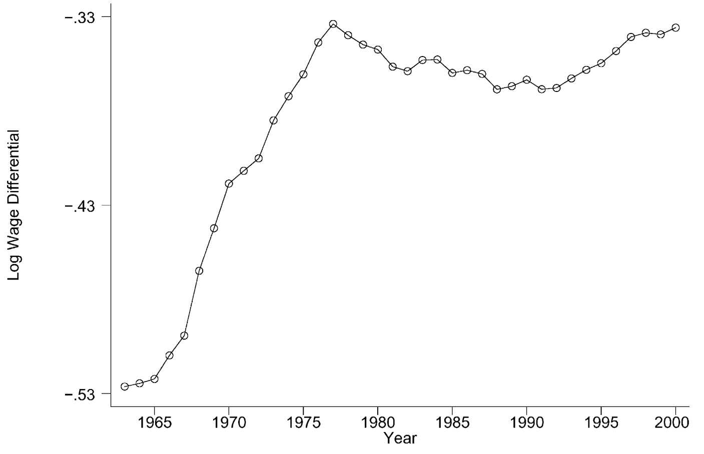

Fig. 1. Raw Earnings Gap of Black Men Relative to White Men, Log Weekly Earnings, 1963–2000.

black and white men employed full-time, full-year (i.e. working at least 35 hours per week in a typical week and employed for at least 50 weeks per year).[8] Over the entire time series, a sizable race gap in the earnings of men is present, never falling below 0.33 on the log scale for weekly earnings (which corresponds to a difference, for example, of approximately 835 dollars per week for whites and only 600 dollars per week for blacks). Even though the black-white gap remains large over the entire time series, Fig. 1 shows that the measured gap has declined, as argued in the literature summarized earlier. However, the initial increase in the gap in the 1980s, followed by its decline thereafter, awaits a compelling explanation. And, it is therefore this last period on which we focus in this article, as it also corresponds to the main period of growth in earnings inequality across the entire labor market.

Figure 2 presents five-year moving averages of the weekly earnings of all white and black men employed full-time, full-year (hereafter, FTFY) in classes I, II, IIIa, IIIb, V, VI, and VIIa from 1982 to 2000. The well-documented growth in earnings inequality is evident, although here for the first time we can see it as growth in between-class inequality. In particular, the class-specific trend lines diverge from 1982 onward and do not narrow substantially even in the robust economic expansion of the late 1990s that lifted the wages of all workers. The weekly earnings of workers in class V, which includes supervisors of manual workers, fell below the average wages of workers in class II. Likewise, the weekly earnings of semi-skilled and unskilled manufacturing workers in classes VI and VIIa fell below those of

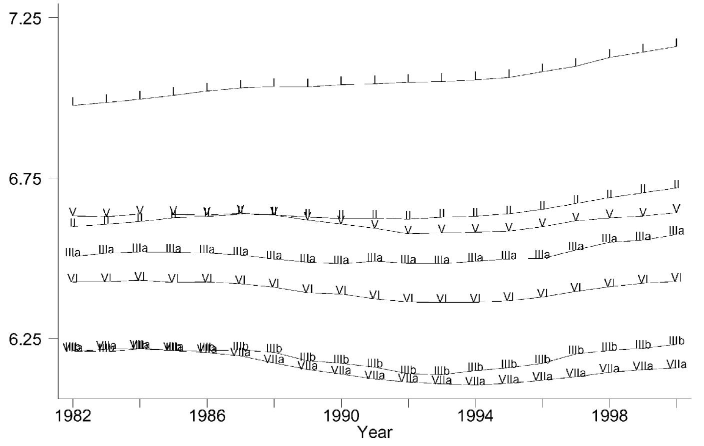

Fig. 2. Log Weekly Earnings by EGP Social Class for Black and White Males Who Work Full-time, Full-Year.

routine non-manual workers in classes IIIa and IIIb. All three trends reflect the relative decline in the fortunes of those employed in the manufacturing sector.

To examine how changes in the black-white gap are associated with general increases in inequality across the labor market, the first step is to examine changes in the distributions of white and black men across social classes. Accordingly, we analyze pooled data from the last five years of the two most recent economic expansions in the United Sates, 1985 through 1989 and 1996 through 2000.[9]

For each of these two time periods, Table 2 presents the proportion of FTFY workers employed in each of the EGP classes summarized in Table 1 separately for white and black men. Two trends are clear. First, the black-white difference in the distribution of men across social classes declined from the late 1980s through the 1990s. Sizable differences still remain, with blacks much less likely to be employed in classes I and II, but their proportional representation within classes continued to converge to that of whites. Second, the class distribution among white and black men has changed slightly, and this structural shift is most clearly revealed in over-time differences in the proportions of white men in each of the EGP classes (since the changes over time for blacks are confounded by inter-generational upward mobility). In particular, the proportion of whites who are employed in classes I and VIIa increased while the proportion employed in classes II and V decreased.

As shown in Fig. 2, the earnings advantages that classes I, II, IIIa, and IIIb have over classes V, VI, and VIIa increased in the 1980s and 1990s. Even though black men moved out of classes VI and VIIa over this time period, the declining relative earnings of these classes, in which blacks were still much more likely

Table 2. Proportion of Full-time, Full-year White and Black Males in Each Social Class for the Last Five Years of the Two Most Recent Economic Expansions.

	White Males (Ages: 18–64)		Black Males (Ages: 18–64)	
	1985–1989	1996–2000	1985–1989	1996–2000
I	0.193	0.199	0.082	0.097
II	0.236	0.227	0.137	0.163
IIIa	0.061	0.059	0.086	0.082
IIIb	0.022	0.025	0.019	0.031
IVab	0.035	0.032	0.014	0.019
IVc	0.009	0.008	0.0003	0.001
V	0.076	0.067	0.067	0.068
VI	0.158	0.156	0.136	0.128
VIIa	0.197	0.213	0.449	0.407
VIIb	0.011	0.014	0.010	0.005

to be employed, limited the progress that otherwise would have been achieved in mitigating the unconditional black-white gap. We therefore conclude that the basic social class decomposition presented jointly in Fig. 2 and Table 2 is consistent with the dominant interpretation of economists through the mid-1990s (see the discussion of Altonji & Blank, 1999, earlier). Blacks may have experienced increasing educational and occupational attainment, but declines in the overall black-white gap in earnings were modest at best because of general increases in inequality.

We now look within social classes at the black-white gap in order to understand how the gap varies across social classes and, more importantly, how this variation has changed over time. The social class breakdown allows for a theoretically-grounded exploration of the heterogeneity of the black-white gap across different locations in the labor market. Of course, there is considerable controversy over whether the EGP schema locates theoretically meaningful categories of positions (see debates cited earlier), centering on whether one is willing to accept the general employment relations perspective of Erikson and Goldthorpe (1992), most recently elaborated in Goldthorpe (2000). Although we do not wish to strongly endorse such arguments, we would argue that: (1) they have not been refuted by anyone; and (2) at least for our project, as we will now show, the EGP schema does generate a pattern of interpretable heterogeneity of the black-white gap that we see as compelling.

Accordingly, for Fig. 3, we calculated the raw wage gap within each of the EGP classes analyzed for Fig. 2, again only for black and white male FTFY workers.

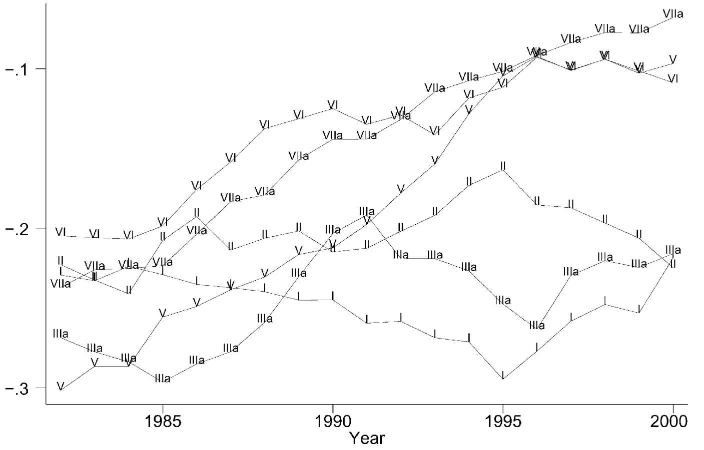

Fig. 3. Black-White Earnings Gap by Social Class.

To be precise, the estimates plotted in Fig. 3 are five-year moving averages within each EGP class of a simple difference in sample expectations:

$$E_N[Y_i|B_i = 0, \mathrm{FTFY}_i = 1] - E_N[Y_i|B_i = 1, \mathrm{FTFY}_i = 1], \qquad \text{(E1)}$$

where FTFY_i is a dummy variable for full-time, full-year workers.[10]

Figure 3 demonstrates that there is substantial heterogeneity across social class of the black-white gap in the weekly earnings of men. In particular, the black-white gap seems to have narrowed dramatically for class V (supervisors of manual workers and lower grade technicians) and to a lesser but still substantial degree for classes VI and VIIa (skilled, semi-skilled, and unskilled manual workers). At the same time, however, the gap may have increased slightly for class I (higher-grade professionals and managers) and decreased only slightly for class II (lower-grade professionals and managers; higher-grade technicians; supervisors of non-manual workers) and class IIIa (higher-grade non-manual workers).[11]

If this pattern is genuine, then it has important implications for explanations of both the black-white gap in the earnings of men and general increases in earnings inequality across classes. Before laying out these implications, in the next section we will show that the trends are, as best we can tell, genuine. They are substantively non-trivial, statistically significant, cannot be explained by trends in the observable characteristics of black and white men, and are not produced by self-selection out of employment because of race-specific costs or other institutional factors.

A Formal Linear Test of the Relative Decline in the Black-White Gap

Although results such as those presented in Figs 1 through 3 can be regarded as primarily descriptive, when enacting adjustments for observed and unobserved characteristics, one inevitably confronts the challenges of causal inference. In an attempt to be as explicit as possible about how we will interpret the results of this section, we will adopt the counterfactual model of causality (see Winship & Morgan, 1999 for a review).[12] Our basic strategy is to form estimates of the potential earnings of blacks if they were instead white and then average these differences over the distribution of: (1) blacks who are employed full-time, full-year; or (2) the predicted distribution of blacks who would be employed full-time, full-year if they were as likely as whites to be employed full-time, full-year. We denote the potential earnings of each individual if black as Y_i^{b} and the potential earnings of each individual if white as Y_i^{w}. In the CPS data we analyze, we have at our disposal individual realizations of two related but substantially more limited random variables: Y_i, an individual's observed earnings and B_i, a dummy variable equal to 1 if a sampled individual self-identifies as black. The most general gap we will consider is:

$$E[\gamma_{i\in \mathrm{B}}] \equiv E[Y^{\mathrm{w}}_{i\in \mathrm{B}}] - E[Y^{\mathrm{b}}_{i\in \mathrm{B}}] \tag{G1}$$

where B denotes the subset of the population that, if asked, would self-identify as black. Gap G1 is explicitly defined as the average black-white gap among blacks, and, as a result, it focuses attention on the earnings penalties associated with being black. Accordingly, we will ignore the earnings penalties whites would suffer if they were suddenly found to be black. And, we will not attempt to estimate the black-white gap across all white and black men, $E[Y_i^{\mathrm{w}} - Y_i^{\mathrm{b}}]$, as it is not clear that (1) it is at all practical to do so or (2) that there are any analytic benefits to modeling this more encompassing gap.

We maintain two types of conditions when estimating and then interpreting our adjusted results. First, the contrasts we will attempt to estimate will be very narrow (indeed, as we will show below when defining gaps G2 and G3, more narrow than even gap G1). The baseline race contrast we will attempt to estimate is the average earnings gain an employed black male working full-time, full-year would expect to capture if he were suddenly found to be white. When we move beyond this contrast, we do so only very minimally, attempting to estimate the gap among a set of hypothetically employed blacks who would be employed full-time, full-year if they were as likely as whites to have access to stable employment in the core full-time, full-year labor market.

We will not proceed any further in this regard, such as by trying to estimate what the black-white gap would be if black men had grown up in families with as much family income as whites, and hence would have had the available resources with which to purchase higher quality secondary and post-secondary education. If we were to attempt to statistically equalize all possible pre-market characteristics (see Neal & Johnson, 1996) across white and black men (i.e. switching to the contrast "what if a black male were to have been born white and into a white family"), we would need to assert additional assumptions on top of those we assert below, clouding the basic labor market contrasts that we regard as the subject of primary theoretical interest in this study. Attempting to estimate the average gains that black labor market participants would capture if they were subject to race-neutral hiring and compensation policies is challenging enough already.

Second, we stipulate that our estimates inform us only about potential wage gains that a few blacks might capture if the equilibrium of the labor market remained the same. That is, because we cannot infer how the entire distribution of earnings would be altered if, in theory, all employers were simultaneously to become race-blind, the estimates we will offer are only informative about the expected gain that a few black men could expect to receive if they were suddenly found to be white by a few randomly chosen employers.[13]

Adjustments for Observed Characteristics

Table 3 reports results from a formal linear test of the relative decline in the black-white gap for classes V, VI, and VIIa between 1982 and 2000. We fit nine variants of the model:

$$\begin{aligned} Y_i = \mathbf{Z}_i\beta + \tau_i[B_i \times T_i] + \delta_{II}[B_i \times T_i \times II_i] + \delta_{IIIa}[B_i \times T_i \times IIIa_i] + \cdots \\ + \delta_{VIIa}[B_i \times T_i \times VIIa_i] + \varepsilon_i \end{aligned} \tag{1}$$

where B_i is a dummy variable for blacks, T_i is a linear term for time (in years), and II_i through $VIIa_i$ are dummy variables for classes II through VIIa (and hence, where class I is the omitted reference class). For all models in Table 3, the covariate vector $\mathbf{Z}_i$ includes a constant, main effects for race, time, and social class (B_i, T_i, and II_i through $VIIa_i$) as well as two-way interactions between race and class ($B_i \times II_i$ through $B_i \times VIIa_i$). In the second panel, main effects for basic human capital and demographic characteristics are added to $\mathbf{Z}_i$: years of education, years of potential experience and its square, three region dummies, two dummies for urbanicity, and four dummies for types of marital status. In the third panel, two-way interactions are included between each of the class dummies and the additional variables for human capital and demographic characteristics.

Within each panel, three separate sets of weights are used for the regressions. For the first column of each panel, the weights are the individual sampling probabilities of being included in the CPS, standardized by the relative size of

Table 3. Variations of a Linear Test for a Greater Relative Decline in the Race Gap in Classes V, VI, and VIIa.

	No Covariates			Human Capital and Demographic Covariates			Human Capital and Demographic Covariates Interacted with Class		
		Weights 1985–1989	Weights 1996–2000		Weights 1985–1989	Weights 1996–2000		Weights 1985–1989	Weights 1996–2000
τ_{I}	−0.123	0.020	−0.005	0.086	0.025	−0.013	0.135	0.065	0.032
	(0.207)	(0.192)	(0.117)	(0.187)	(0.178)	(0.121)	(0.193)	(0.183)	(0.134)
δ_{II}	0.371	0.487	0.481	0.261	0.419	0.432	0.205	0.402	0.413
	(0.364)	(0.249)	(0.222)	(0.335)	(0.249)	(0.218)	(0.341)	(0.264)	(0.235)
δ_{IIIa}	0.503	0.803	0.658	0.240	0.586	0.572	0.136	0.525	0.502
	(0.279)	(0.302)	(0.193)	(0.234)	(0.248)	(0.193)	(0.253)	(0.247)	(0.236)
δ_{IIIb}	0.689	0.672	0.551	0.337	0.140	0.024	0.307	0.248	0.119
	(0.447)	(0.694)	(0.635)	(0.427)	(0.499)	(0.433)	(0.406)	(0.494)	(0.441)
δ_{V}	1.492	1.225	1.168	1.172	0.934	0.895	1.142	0.963	0.927
	(0.330)	(0.248)	(0.214)	(0.289)	(0.301)	(0.247)	(0.289)	(0.305)	(0.254)
δ_{VI}	0.751	0.836	0.959	0.509	0.599	0.744	0.456	0.592	0.735
	(0.290)	(0.243)	(0.204)	(0.298)	(0.228)	(0.190)	(0.304)	(0.229)	(0.198)
δ_{VIIa}	1.168	0.857	0.886	0.597	0.543	0.565	0.576	0.529	0.559
	(0.214)	(0.243)	(0.164)	(0.189)	(0.207)	(0.135)	(0.186)	(0.195)	(0.136)
Average decline in the race-gap within classes V, VI, and VIIa	1.013	0.993	0.999	0.845	0.692	0.721	0.860	0.760	0.772
	(0.084)	(0.099)	(0.098)	(0.087)	(0.227)	(0.001)	(0.083)	(0.001)	(0.001)
Average decline in classes V, VI, and VIIa relative to classes I, II, IIIa, and IIIb	0.746	0.482	0.582	0.549	0.406	0.478	0.563	0.401	0.482
	(0.126)	(0.216)	(0.174)	(0.001)	(0.228)	(0.002)	(0.001)	(0.221)	(0.002)
p-value for relative decline	<0.001	0.038	0.004	<0.001	0.092	0.010	<0.001	0.086	0.010

Note: Standard error are in parentheses and are robust, heteroskedastic-consistent standard errors, further adjusted for the clustering of respondents within CPS years. All coefficients and standards errors are multiplied by 100. The average decline is stipulated to be the linear combination of coefficients: $[(\tau_I + \delta_V) + (\tau_I + \delta_{VI}) + (\tau_I + \delta_{VIIa})]/3$, and the relative decline is stipulated to be the linear combination of coefficients: $[(\tau_I + \delta_V) + (\tau_I + \delta_{VI}) + (\tau_I + \delta_{VIIa})]/3 - [\tau_I + (\tau_I + \delta_{II}) + (\tau_I + \delta_{IIIa}) + (\tau_I + \delta_{IIIb})]/4$. The *p*-value is for a two-tailed test where the null hypothesis is that the relative decline is 0. Within each panel, the *N*s for the models in the first through third columns are 450,982, 444,554, and 443,339.

the CPS sample in each year. For the second and third columns within each panel, these standardized sampling weights are multiplied by additional scaling factors in order to generate estimates of average wage gaps while aligning the within-class occupational distributions of both whites and blacks in every year to the same unit-level occupational distribution of blacks for two alternative reference time period periods, 1985–1989 and 1996–2000. These modeling weights were generated by multiplying the sampling weights of each individual by additional occupation- and race-specific yearly scaling factors:

$$\pi^{\mathrm{w}}(k, t) = \frac{\text{\% of black FTFY workers in occupation } k \text{ in the reference time period}}{\text{\% of white FTFY workers in occupation } k \text{ in year } t}$$

for white males and

$$\pi^{\mathrm{b}}(k, t) = \frac{\text{\% of black FTFY workers in occupation } k \text{ in the reference time period}}{\text{\% of black FTFY workers in occupation } k \text{ in year } t}$$

for black males.[14] When deployed in weighted OLS models, these weights balance unit-occupation-level differences across race, and as a result, adjust for within-class race differences in occupational attainment. At the same time, they sweep away heterogeneity bias of the form that Grusky and Weeden (2001) claim weakens the claims of all big class models. And, moreover, they do so not simply by pushing the heterogeneity aside, but instead by normalizing it across race in the accordance with the definition of gap G1 and with reference to two separate and theoretically meaningful time periods.

The model reported in the first column of Table 3 does not include any human capital or demographic covariates. As a result, the reported coefficient estimate of τ_{I} represents the linearly interpolated slope of the line for class I in Fig. 3. Similarly, the coefficient estimates of δ_{II} through δ_{VIIa} represent the differences in the linearly interpolated relative slopes of the lines in Fig. 3 for classes II through VIIa in comparison with class I. For economy of space, each of the estimated coefficients and standard errors is multiplied by 100.

The coefficients reported in the first column imply that over the nineteen year time period, the gap for class I increased slightly. The coefficient for τ_I reported in the first row is -0.123. When divided by one hundred and multiplied by nineteen, the coefficient implies that the black-white gap increased by 0.0234 on the log scale. In contrast, on average for classes V through VIIa, the black-white gap declined by 0.192 on the log scale (which is the linear combination of coefficients, $[(\tau_I + \delta_V) + (\tau_I + \delta_{VI}) + (\tau_I + \delta_{VIIa})]/3$, reported in the third row from the bottom of the table as 1.013, and which was then divided by one hundred and multiplied by nineteen). Picking reasonable values for the initial black-white gap, and converting to a dollar metric over 52 weeks of work, the model in the first column implies that the black-white gap for class I increased from 12,614 dollars to 13,641 dollars whereas the black-white gap averaged over classes V, VI, and VIIa declined from 7,650 dollars to 1,949 dollars.[15]

Although this contrast compares the trend in the black-white gap within class I to those within classes V, VI, and VIIa, it ignores classes II, IIIa, and IIIb, and hence is perhaps too narrow of a test. As shown in Fig. 2, the earnings of workers in classes V, VI, and VIIa generally fell behind those of workers in classes I, II, IIIa, and IIIb. If the overall increase in earnings inequality is driving this change, we should expect a relatively greater decline in V, VI, and VIIa in comparison to all other classes. The second to last row of the table reports the average class-specific linear trend in the black-white gap for classes V, VI, and VIIa relative to the average class-specific linear trend in the black-white gap for classes I, II, IIIa, and IIIb:

$$\frac{(\tau_I + \delta_V) + (\tau_I + \delta_{VI}) + (\tau_I + \delta_{VIIa})}{3} - \frac{\tau_I + (\tau_I + \delta_{II}) + (\tau_I + \delta_{IIIa}) + (\tau_I + \delta_{IIIb})}{4}$$

along with its corresponding standard error (again where both are multiplied by 100 for economy of space). The last row of the table then reports the *p*-value for a two-tailed *t*-test, where the null hypothesis is that this linear combination of coefficients is 0. According to frequentist standards, if there were no average class difference in the decline in the black-white gap between classes I, II, IIIa, and IIIb relative to classes V, VI, and VIIa (under the specification of the model in column 1), the probability that we would observe a relative difference between classes as large in absolute value as is reported in column 1 is less than 0.001.

For the remaining eight models reported in Table 3, the same general pattern holds. The unreported main effect coefficients in $\mathbf{Z}_i$ change substantially when covariates are added, such that overall class differences in earnings are partialled out across both class and educational training. But the overall within-class trends in the black-white gap follow the same pattern, as shown in the rows for the average decline in the gap for classes V, VI, and VIIa.

Additional Adjustments for Selection Patterns into the Labor Force

Underlying labor force participation patterns vary by race, complicating the development of a consistent estimator for the black-white gap. To focus on the essential problem, consider the estimation challenges introduced by the following self-selection narrative. According to the literature on urban labor markets, a spatial mismatch between jobs and the residence locations of blacks generates greater relative costs of employment for blacks (e.g. greater transportation costs;

see Holzer, 1991). As a result of this difference in costs, blacks at the bottom of the distribution of potential occupational attainment may have higher wage thresholds above which they would choose to work rather than remain idle. Accordingly, blacks who would receive the lowest wages in the lowest-paying occupations are less likely to enter employment than comparable whites, and hence those blacks who are observed in low-paying occupations receive relatively higher wages on average than their unemployed counterparts would necessarily receive if they chose to work. This line of argument suggests that a smaller within-occupation earnings gap is observed at the lower end of the occupational distribution because the average wages of blacks in classes VI and VIIa are artificially inflated by self-selection out of employment.[16]

As fully described in the appendix, we assess the robustness of our conclusions based on Fig. 3 and Table 3 to these types of selection bias by first stipulating that the gap of primary theoretical interest is the gap at the core of the labor market for FTFY workers. We then perform an adjustment for selection patterns by artificially equalizing across race the proportion of workers employed FTFY and then calculating the gap assuming that the additional black workers who are added to the "what if" FTFY sample had the lower weekly earnings typical of full-time, part-year (hereafter, FTPY) workers in the same social class.

Our implementation of this strategy follows several steps. First, for each class c in each year t, we calculated the proportion of observed white and black FTFY and FTPY workers who were FTFY workers:

$$\lambda^{\mathrm{w}}(c,t) = \frac{\text{\% of white FTFY workers in class } c \text{ in year } t}{\text{\% of white FTFY and FTPY workers in class } c \text{ in year } t}$$

and:

$$\lambda^{\mathrm{b}}(c,t) = \frac{\text{\% of black FTFY workers in class } c \text{ in year } t}{\text{\% of black FTFY and FTPY workers in class } c \text{ in year } t}$$

We then used the CPS data (along with external data on the yearly incarceration rates of the black and white male adult population) in order to estimate a broad-based rate of FTFY employment:

$$\theta^{\mathrm{b}}(c,t) = \frac{\text{\% of black FTFY workers observed in class } c \text{ in year } t}{\begin{array}{c}\text{\% of all blacks who would be in class } c \text{ in year } t \text{ if they were as likely} \\ \text{as whites to be employed as FTFY workers in class } c \text{ in year } t\end{array}}$$

where we invoked a reasonable (and pessimistic) assumed distribution for the employment of additional blacks across the class distribution. Values for these three rates, along with the specific components on which they are based, are presented as averages over the 1985–1989 and 1996–2000 time periods in appendix Tables A1 and A2.

For execution of the selection adjustment models, we then added all FTPY black males to the analysis sample and reestimated the models in Table 3 with the augmented sample. While doing so, we multiplied the model-specific weights used for the nine columns of Table 3 by additional year-specific scaling factors for black workers only. For FTFY blacks, the weights were multiplied by:

$$\psi^{b}(c,t)=\frac{\theta^{b}(c,t)}{\lambda^{w}(c,t)},$$

and for FTPY blacks, the weights were multiplied by:

$$\psi^{b}(c,t)=\left(\frac{\lambda^{b}(c,t)}{1-\lambda^{b}(c,t)}\right)\left(\frac{\lambda^{w}(c,t)-\theta^{b}(c,t)}{\lambda^{w}(c,t)}\right).$$

In combination, the modified weights for FTFY and FTPY blacks lower the relative representation of FTFY blacks and raise the relative representation of FTPY blacks. And since FTPY blacks have earnings that are approximately 20% less than FTFY blacks (see columns eight through twelve of Table A2), the modified weights have the effect of lowering the average earnings of blacks by an amount that is proportional to the difference between observed white and black FTFY participation rates in each year.

The resulting models are presented in Table 4. Comparing the first column of this table with that of Table 3, the coefficient for the relative average decline is reduced from 0.746 to 0.699. Thus, only approximately six percent of the baseline relative decline can be attributed to trends in race differences in selection into the FTFY labor force. A similar pattern is revealed in the remaining columns of the table, using covariance adjustments alongside the weighting estimators. In the last six columns, however, the effect of the selection adjustment is mitigated by the human capital covariates; FTPY blacks have substantially lower educational attainment, as do all other categories of the "what if" blacks.[17]

Selection into the labor force is an important consideration, and the unreported main effects show that the relative size of the wage gap within classes V, VI, and VIIa is indeed substantially reduced when the sample is augmented with FTPY black workers and then weighted according to our selection-adjustment scheme. But, this does not appear to explain away the time trend that is revealed in Fig. 3 and Table 3. As shown in the appendix, between the 1980s and 1990s, the trend toward greater incarceration and "voluntary" non-participation of blacks seems roughly counterbalanced by shifts out of unemployment and relative shifts out of part-year work. Thus, even though black males remained substantially less likely to be employed FTFY by the end of the time series, the race difference in the rate of FTFY employment changed less than might be expected.

Table 4. Selection-Adjusted Linear Tests for a Greater Relative Decline in the Race Gap in Classes V, VI, and VIIa.

	No Covariates			Human Capital and Demographic Covariates			Human Capital and Demographic Covariates Interacted with Class		
		Weights 1985–1989	Weights 1996–2000		Weights 1985–1989	Weights 1996–2000		Weights 1985–1989	Weights 1996–2000
τ_1	−0.192	−0.078	−0.063	−0.055	−0.037	−0.034	−0.007	0.015	−0.007
	(0.206)	(0.192)	(0.123)	(0.198)	(0.195)	(0.155)	(0.204)	(0.217)	(0.177)
δ_{II}	0.401	0.517	0.538	0.335	0.531	0.495	0.317	0.503	0.495
	(0.340)	(0.326)	(0.325)	(0.329)	(0.337)	(0.333)	(0.326)	(0.351)	(0.343)
δ_{IIIa}	0.416	0.806	0.635	0.215	0.653	0.547	0.158	0.593	0.475
	(0.331)	(0.360)	(0.283)	(0.256)	(0.285)	(0.229)	(0.276)	(0.304)	(0.243)
δ_{IIIb}	1.071	0.898	0.782	0.548	0.327	0.112	0.611	0.446	0.210
	(0.381)	(0.519)	(0.574)	(0.375)	(0.416)	(0.439)	(0.384)	(0.406)	(0.426)
δ_V	1.571	1.306	1.059	1.250	1.033	0.826	1.312	1.081	0.892
	(0.312)	(0.287)	(0.170)	(0.299)	(0.272)	(0.199)	(0.302)	(0.288)	(0.218)
δ_{VI}	0.847	0.905	0.973	0.611	0.715	0.809	0.604	0.692	0.823
	(0.270)	(0.239)	(0.209)	(0.294)	(0.259)	(0.235)	(0.304)	(0.279)	(0.260)
δ_{VIIa}	1.095	0.779	0.749	0.626	0.527	0.529	0.637	0.525	0.579
	(0.217)	(0.238)	(0.148)	(0.189)	(0.212)	(0.149)	(0.187)	(0.219)	(0.168)
Average decline in the race-gap within classes V, VI, and VIIa	0.979	0.919	0.864	0.774	0.721	0.687	0.843	0.781	0.758
	(0.071)	(0.090)	(0.018)	(0.077)	(0.094)	(0.111)	(0.076)	(0.092)	(0.106)
Average decline in classes V, VI, and VIIa relative to classes I, II, IIIa, and IIIb	0.699	0.441	0.439	0.554	0.381	0.433	0.579	0.381	0.469
	(0.001)	(0.178)	(0.168)	(0.104)	(0.165)	(0.151)	(0.001)	(0.154)	(0.144)
p-value for relative decline	<0.001	0.023	0.018	<0.001	0.033	0.010	<0.001	0.024	0.004

Note: See Table 3. Within each panel, the *N*s for the models in the first through third columns are 459,321, 454,109, and 451,505.

Although more research is surely needed on the consequences of selection into the labor force, the decline in the wages of white males appears to be driving the trends we have presented, not artefactual increases in the wages of black males. Taken together, the more refined results we have presented in Tables 3 and 4 do not change the basic result revealed in Fig. 3. The black-white gap among men has declined relatively more at the bottom of the class distribution.

DISCUSSION

Between 1982 and 2000, the black-white gap in the earnings of men decreased slightly while aggregate social class differences in earnings increased. The black-white gap for class I is either constant or slightly increasing, and yet it emerges as the largest gap across all classes primarily because the relative wages of white men from classes V, VI, and VIIa have fallen consistently over the past two decades. In the context of growing wage inequality in general, white men at the bottom of the class distribution have been unable to maintain their wage advantages over their black counterparts to the same degree that white men at the top of the class distribution have. In this concluding section, we will interpret the relative decline as consistent with the rent destruction explanation for increases in earnings inequality developed by Sørensen (1996, 2000). However, before proceeding to a discussion of rent destruction, and its implications for research on social class and social stratification in post-industrial societies, we first discuss other plausible explanations for the findings we present.

Kaufman (1983) argued that gains in the average occupational placement of blacks would generate a paradoxical increase in the net wage gap between white and black men, since he found when analyzing 1970 census data that the race gap was slightly larger at the top of the distribution of occupations (after a variety of covariance adjustments). Grodsky and Pager (2001) then analyzed 1990 census data to show that in spite of demonstrable gains in the occupational attainment of black men, the black-white gap in earnings persists and also seems to remain largest at the top of the distribution of occupational prestige and average earnings. These findings are consistent with Kaufman's prediction, even though they do not explicitly demonstrate an increase in the gap. In addition, Grodsky and Pager supplement Kaufman's explanatory narrative, offering a new explanation for the pattern of heterogeneity across the occupational distribution. They argue that white professionals serve clients that on average have a greater capacity to pay high fees for their services. Whites in these occupations occupy structural job niches that command greater profitability than do blacks. Although we have no direct evidence to support Grodsky and Pager's explanation, we are quite captivated by its potential, especially since it is complementary to and consistent with the rent destruction interpretation we adopt below.[18]

At the same time, two other explanations could generate our findings (and those of Grodsky & Pager as well). First, as detailed earlier, patterns of selection out of the labor force could generate the relative decline in the race gap, even though our analysis suggests otherwise. Second, a relatively greater decline in discriminatory compensation practices at the bottom of the class distribution could also give rise to the same empirical pattern. Again, however, it is not clear that a relative decline in discriminatory compensation practices should be interpreted as an alternative to rent destruction, rather than as part and parcel of the same mechanism. To explain why, we now present an explanation that we see as the most compelling and theoretically intriguing, even though we recognize that we have little direct evidence to favor it over these alternatives.

Sørensen (1996, 2000) argued that when examining labor market processes, stratification researchers should be interested in the distribution of three different quantities: (1) Y^a, actual wages paid in the labor market; (2) Y^c, wages that would be paid under perfect competition; and (3) $r^c = Y^a - Y^c$, rent. Adopting this simple but broad framework, Sørensen argued that increases in labor market inequality in the U.S. reflect pervasive but selective rent destruction since the 1970s.

Even though Sørensen offered little detail on precisely how this rent destruction occurred, one can develop a set of implied claims using the EGP class schema. Since the 1970s, individuals in the most privileged jobs within classes I and II have been able to retain (or even increase) their wage advantages over others in less privileged jobs (both within classes I and II and in comparison with all jobs in other social classes). In contrast, in classes V, VI, and VIIa, the wage advantages associated with the best jobs in the 1960s and 1970s have declined so that all wages paid to workers within classes V, VI, and VIIa have converged to wage levels for the least paid.

Sørensen's argument is more subtle than claims about the hollowing out of the class distribution in post-industrial society, since the motivation for the framework is to transcend the simple story of how the decline in the prevalence of jobs in heavy industry produced a new market equilibrium in which an over-supply of unskilled workers reduced the average wage levels of classes V, VI, and VIIa. Rather, according to Sørensen, the recent increases in earnings inequality reflect a fundamental shift in employment relations, organized around new norms for firm management of individuals' careers. Although workers at the top of the class distribution continue to be paid according to lifetime earnings schemes, which often include loyalty and efficiency wage components, workers at the bottom of the class distribution have become increasingly likely to be paid according to managers' calculations of workers' instantaneous marginal contributions to output.

Our finding of a relative decline in the black-white earnings gap among men within classes V, VI, and VIIa is consistent with Sørensen's rent destruction explanation *if* one is willing to assert assumptions about race differences in access to rent-advantaged positions. For example, relying on the classic split-labor market argument of Bonacich (1972, 1976), or, alternatively, the characterization of black-white differences in Sørensen (1975, 1979), one could assume that prior to 1980 black men were more likely to be excluded from rent-advantaged positions across the entire spectrum of available occupations, and that, after 1980, access to such rent-advantaged positions increased no more for those in classes V, VI, and VIIa than for classes I and II (and perhaps IIIa and IIIb). If selective rent destruction is a plausible explanation for increases in inequality between social classes after 1980, then the destruction of rent-advantaged positions should be greatest at the bottom of the class distribution, and a *ceteris paribus* implication follows immediately from the assumption above: The black-white gap in earnings among men should have decreased relatively more within classes V, VI, and VIIa than within classes I and II (and perhaps IIIa and IIIb).[19] This is precisely the pattern revealed in the analysis we report.

We therefore conclude that the rent destruction hypothesis promoted by Sørensen (1996, 2000) is a useful explanatory narrative for the specific set of findings that we present, and one which can be used to supplement rather than supplant first-order interpretations suggested by the EGP-class-based analysis on its own. A larger implication of our analysis, if this interpretation is convincing, is that class analysis in its traditional modes is not irrelevant and need not be completely overhauled to remain useful. When big classes based on reasonable codings of census occupational schemes are deployed in analysis, a compelling descriptive picture of labor market outcomes can be generated. Thereafter, if it is so desired, supplementary explanatory effort can be mounted to elucidate the consequences of unappreciated heterogeneity, as argued by Grusky and Weeden (2001), or more basic structural processes, as argued by Sørensen (1996, 2000).

NOTES

1. In this article, we will not investigate gender differentials or contrasts with racial/ethnic groups other than those identified for the Current Population Surveys as white and black. These differentials are worthy of analysis as well. However, the gap between white and black men has been the subject of much targeted research and represents, we would argue, the best starting point for developing an EGP-class-based analysis of changes in earnings inequality.

2. Hout et al. (1993) is the only published class-based decomposition of earnings inequality in the U.S. that we have found which uses the EGP class schema (which

is what we regard as the dominant class schema employed in empirical research; see below).

3. The technological explanation has not gone completely unexamined in sociology, as Fernandez (2001) provides a fascinating account of how technological innovation in one plant resulted in the skill-upgrading of occupations associated with that plant. Whether or not this study "provides strong evidence for the importance of skill-biased technological change as an explanation of rising wage inequality" (Fernandez, 2001, p. 316) is a matter for debate, one which should properly rest on the logic of Fernandez's claim that his study's "natural-experiment design solves the self-selection problem that plagued past skill-bias studies" (Fernandez, 2001, p. 315).

4. Sociologists have offered explanations for persistent race differences in labor market outcomes, focusing on the ways in which discriminatory hiring practices are institutionalized in job creation and staffing decisions, grounded either in the dual labor market conceptualizations of Bonacich (1972, 1976) or the spatial mismatch hypotheses of Wilson (1987). Complementing these structural explanations, attention has also focused on differentials in opportunities to acquire marketable skills and credentials prior to entering the labor market (e.g. Farley, 1984), differences in cultural competencies that are rewarded in the labor market (see Lamont, 1999), and the systematically incorrect beliefs that employers maintain about the skills of non-white workers (see Moss & Tilly, 2001). Our claim is only that these explanations have not been deployed by sociologists in explicit attempts to model change in the past two decades in black-white differences in earnings.

5. In this sense, it is not especially problematic, from our perspective, that we leave out of our descriptive picture all women and those men not identified as white or black for the CPS data. Clearly, however, if one regards classes as holistic entities, then such truncation of the class structure would be very problematic, especially if one were interested in characterizing comprehensively how the race and gender compositions of social classes are evolving in postindustrial society.

6. Consider the following comparison. For the 450,982 respondents in the first panel of Table 3, we fit a regression equation predicting weekly earnings with parameters for time, experience, experience squared, region, urbanicity, and marital status. The R^2 for this baseline model was 0.1551. Adding to this baseline model linear and quadratic terms for years of education, each interacted with time, the R^2 increased to 0.3281. However, when adding six dummy variables for social class, each interacted with time, to the same baseline model, the R^2 increased to 0.3353. Of course, EGP uses more parameters, but by the criterion of variance explained, it does outperform a standard Mincerian human capital rate of return model, even when augmented with a quadratic term for education.

7. Our procedures for handling topcodes allow us to estimate conditional means, but they do not determine the qualitative conclusions we report to any substantial degree. The models in Tables 3 and 4 were re-estimated using median regression procedures and the results were only trivially different.

8. For comparability with later results, Fig. 1 is further restricted to those in one of the seven main social classes we will analyze throughout the article (i.e. excluding farmers and most self-employed workers; see below).

9. For now, we do not analyze recessionary years (1982, 1990, and 1991) and early recovery years (1983, 1984, and 1992 through 1995) in order to focus analytically on stable full-employment distributions of social classes. In the more formal regression tests, these years are included in analysis.

10. In this article, we use $E[\cdot]$ to denote the population expectation and $E_N[\cdot]$ to denote the sample expectation (i.e. the estimated mean from a sample of size N). This notation, drawn from the method of moments literature, allows for economy of expression when making claims about the large sample properties of estimators.

11. For Fig. 3, we do not present the estimated race gap for class IIIb. There are very few black men who hold such occupations in the CPS (approximately 50–80 per year spread across 20 different occupations), and as a result, estimates of the wage gap vary dramatically from year to year, reflecting little more than sampling error. We do, however, include class IIIb in the more formal tests of heterogeneity reported later.

12. We fully recognize that some researchers maintain that counterfactuals cannot be used to profitably examine non-manipulable discrete variables, such as race and gender (see Holland, 1986). In our opinion, Glymour (1986) provides a cogent defense of using counterfactuals to elucidate effects that are presumed to have been caused by a set of potentially manipulable factors that are merely referred to collectively by nominal labels.

13. We fully recognize that these assumptions are very stringent, and that the use of counterfactual models for examining race differences within the labor market will be controversial for some readers. However, we challenge others to be similarly specific in laying out their assumptions and the causal inferences their results can sustain (or, if not making causal claims, why a descriptive analysis is useful).

14. See Handcock and Morris (1999) for a comprehensive treatment of non-parametric weighting estimators, of which our weighting procedures are one variant. See Barsky et al. (2002) for a similar application.

15. These calculations assume a black-white gap of 0.25 on the log scale at the beginning of the time series for both groups. For class I, the initial gap is assumed to be 7 and 6.75 on the log scale (or 57,025 and 44,411 in dollars over 52 weeks). For classes V through VIIa, the gap is assumed to be 6.5 and 6.25 on the log scale (or 34,587 and 26,937 dollars over 52 weeks). The implied decreased average earnings of blacks in class I is 6.7266 on the log scale (or 43,384 in dollars over 52 weeks), while for classes V through VIIa it is 6.442 on the log scale (or 32,638 in dollars over 52 weeks).

16. This argument implicitly assumes that the wage offers received by all individuals at the upper end of the distribution of potential occupational attainment exceed their reservation wages. Thus, there is no similar tendency for self-selection out of employment among blacks at the top of the distribution of potential occupational attainment.

17. For models reported in the second and third columns of each panel, the occupational alignment scaling factors that are utilized for the non-parametric weighting estimators are altered slightly to:

$$\phi^{w}(k,t) = \frac{\text{\% of black FTFY and FTPY workers in occupation } k \text{ in the reference time period}}{\text{\% of white FTFY workers in occupation } k \text{ in year } t}$$

for white males and

$$\phi^{b}(k,t) = \frac{\text{\% of black FTFY and FTPY workers in occupation } k \text{ in the reference time period}}{\text{\% of black FTFY and FTPY workers in occupation } k \text{ in year } t}$$

for black males.

18. That is, as we discuss below, whites at the top of Grodsky and Pager's prestige distribution have held on to their rents, whereas those at the bottom of the distribution have lost some of theirs.

19. A more formal derivation of this implication could be developed as follows. Suppose that all measurable effects of skill differences have been partialled out of Y^a within each class. Suppressing subscripts for individuals i and time t (for compactness of notation), continue to define rent as $r^c \equiv Y^a - Y^c$. Now suppose that competitive wages and the sizes of rents for those who receive rents are independent of race within class: $E[Y^c|C] \perp B$ and $E[r|r > 0, C] \perp B$. Defining a disproportionate share of whites who occupy rent-advantaged positions as $\rho^{\mathrm{w-b}} \equiv \Pr[r > 0|C, B = 0] - \Pr[r > 0|C, B = 1]$, suppose $\rho^{\mathrm{w-b}}$ is positive in all classes. If $E[Y^a|C = c, B = 0] - E[Y^a|C = c, B = 1]$ is shown to be decreasing in time for class c, then $(\rho^{w-b}|C = c)E[r|r > 0, C = c]$ is decreasing in time, which implies that either $\rho^{\mathrm{w-b}}$ or $E[r|r > 0]$ (or both) is decreasing in time for class c. The crucial assumption, of course, is that $\rho^{\mathrm{w-b}}$ remains positive (or, more specifically in this application, that it remains relatively more positive in classes V, VI, and VIIa than in class I).

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APPENDIX

Details of Model Specification

To estimate gap G1, we could attempt to: (1) take the observed earnings Y_i of each black respondent; (2) subtract these observed earnings from plausible estimates of each black respondent's potential earnings if white, $\hat{Y}_i^{\mathrm{w}}$; and then (3) average the resulting individual differences, $Y_i - \hat{Y}_i^{\mathrm{w}}$, over the distribution of black respondents. Unfortunately, estimating individual-level counterfactual wage levels is a daunting, if not impossible, task. It is customary to instead estimate only the two expectations on the right-hand side of gap G1. The main challenge, however, is that the simple difference in sample expectations:

$$E_N[Y_i|B_i = 0] - E_N[Y_i|B_i = 1] \tag{E0}$$

does not converge in probability to $E[\gamma_{i\in \mathrm{B}}]$ (i.e. gap G1) because

$$E_N[Y_i|B_i = 0] \underset{p}{\rightarrow} E[Y^{\mathrm{w}}_{i\notin \mathrm{B}}] \tag{A1}$$

and[20]

$$E[Y^{\mathrm{w}}_{i\notin \mathrm{B}}] > E[Y^{\mathrm{w}}_{i\in \mathrm{B}}]. \tag{A2}$$

In other words, estimator E0 is inconsistent for gap G1 because blacks differ from observed whites on characteristics, such as educational attainment, that are relevant for labor market earnings. The inconsistency of Estimator E0 cannot be addressed by simply conditioning or stratifying on observed variables in the CPS. To explain more completely why this is the case, we proceed by first more carefully specifying the black-white gap of interest.

Aside from the lack of opportunity to acquire "pre-market" skills, which we do not explicitly model in this article, there are likely two important penalties associated with being a black male: fewer opportunities for obtaining stable work and lower wages when in stable work. Gap G1 conflates these two penalties in unknown ways. To separate these effects, we use analogous race-conditioned potential outcome and observed variables for full-time, full-year work – $FTFY_i^{\mathrm{b}}$, $FTFY_i^{\mathrm{w}}$, and $FTFY_i$ – corresponding to dummy variables for working FTFY if black, working FTFY if instead white, and observed FTFY status. With these additional variables, we can more carefully specify the counterfactuals of primary theoretical interest.

We start with the assumption that blacks would be more likely to be employed FTFY if they were instead white:

$$\Pr[FTFY_{i\in \mathrm{B}}^{\mathrm{w}} = 1] > \Pr[FTFY_{i\in \mathrm{B}}^{\mathrm{b}} = 1]. \tag{A3}$$

This assumption of differential propensity to find stable work suggests two separate black-white gaps of interest:

$$E[\gamma_{i\in \mathrm{B}}|FTFY_{i\in \mathrm{B}}^{\mathrm{b}} = 1] \equiv E[Y_{i\in \mathrm{B}}^{\mathrm{w}}|FTFY_{i\in \mathrm{B}}^{\mathrm{b}} = 1] - E[Y_{i\in \mathrm{B}}^{\mathrm{b}}|FTFY_{i\in \mathrm{B}}^{\mathrm{b}} = 1] \tag{G2}$$

and

$$E[\gamma_{i\in \mathrm{B}}|FTFY_{i\in \mathrm{B}}^{\mathrm{w}} = 1] \equiv E[Y_{i\in \mathrm{B}}^{\mathrm{w}}|FTFY_{i\in \mathrm{B}}^{\mathrm{w}} = 1] - E[Y_{i\in \mathrm{B}}^{\mathrm{b}}|FTFY_{i\in \mathrm{B}}^{\mathrm{w}} = 1] \tag{G3}$$

where gap G2 is the black-white gap among blacks who are typically employed FTFY and where gap G3 is the black-white gap among blacks who would be employed FTFY if they were instead white (but again, assuming that a hypothetical shift for a black respondent from the potential state $FTFY_i^{\mathrm{b}}$ to its alternative $FTFY_i^{\mathrm{w}}$ does not entail any change in educational preparation or any other observed or unobserved covariate).

Now, returning to the main text, is estimator E1 consistent for either gap G2 or G3? No, just as estimator E0 is not consistent for gap G1, neither is estimator E1 consistent for either G2 or G3 because it is reasonable to assume that:

$$E_N[Y_i|B_i = 0, FTFY_i = 1] \underset{p}{\rightarrow} E[Y_{i\notin \mathrm{B}}^{\mathrm{w}}|FTFY_{i\notin \mathrm{B}}^{\mathrm{w}} = 1], \tag{A4}$$

$$E[Y_{i\notin \mathrm{B}}^{\mathrm{w}}|FTFY_{i\notin \mathrm{B}}^{\mathrm{w}} = 1] > E[Y_{i\in \mathrm{B}}^{\mathrm{w}}|FTFY_{i\in \mathrm{B}}^{\mathrm{b}} = 1], \tag{A5}$$

and

$$E[Y_{i\notin \mathrm{B}}^{\mathrm{w}}|FTFY_{i\notin \mathrm{B}}^{\mathrm{w}} = 1] > E[Y_{i\in \mathrm{B}}^{\mathrm{w}}|FTFY_{i\in \mathrm{B}}^{\mathrm{w}} = 1], \tag{A6}$$

The earnings of FTFY whites are (asymptotically) upwardly biased estimates of the FTFY earnings of blacks if they were instead white. But, this is only the simplest possible case, and as discussed informally in the selection-adjustment section in the main text there may be cases where the inequalities in assumptions A5 and A6 are entirely misleading.

Non-parametric Weighting Estimators

The estimator we introduce in this section (and deploy in columns 2 and 3 of each panel of Tables 3 and 4) allows for the explicit formulation of answers to useful counterfactual questions, such as: How would the earnings gap for each EGP class have evolved between 1982 and 2000 if both whites and blacks had the distribution across occupations that resulted for blacks in the final stages of the last two economic expansions in the United States?

The non-parametric estimator we deploy can be specified in general form as class and year-specific variants of:

$$E_N[\pi^{\mathrm{w}}(k,t)Y_i|B_i = 0, FTFY_i = 1] - E_N[\pi^{\mathrm{b}}(k,t)Y_i|B_i = 1, FTFY_i = 1] \quad \text{(E2)}$$

where the weights $\pi^{\mathrm{w}}(k, t)$ and $\pi^{\mathrm{b}}(k, t)$ are as specified in the main text. Estimator E2 provides a more reasonable estimator for G2 than E1, in the sense that the asymptotic bias of E2 is almost surely lower than for E1. Moreover, as shown in Tables 3 and 4, this non-parametric estimator can be deployed within a regression set-up, which allows for simultaneous covariance adjustments.

Adjustments for Selection Patterns

When conditioning on class, or any other stratifying variable S, the inequalities in assumptions A5 and A6 may not hold if there are complex selection patterns that generate the data. In particular, even though it may be reasonable to assume that:

$$E[Y^{\mathrm{w}}_{i\in\mathrm{B}}|FTFY^{\mathrm{b}}_{i\in\mathrm{B}} = 1, S_i = s] > E[Y^{\mathrm{b}}_{i\in\mathrm{B}}|FTFY^{\mathrm{b}}_{i\in\mathrm{B}} = 1, S_i = s] \quad \text{(A7a)}$$

and

$$E[Y^{\mathrm{w}}_{i\in\mathrm{B}}|FTFY^{\mathrm{w}}_{i\in\mathrm{B}} = 1, S_i = s] > E[Y^{\mathrm{b}}_{i\in\mathrm{B}}|FTFY^{\mathrm{w}}_{i\in\mathrm{B}} = 1, S_i = s], \quad \text{(A7b)}$$

it may also be the case that:

$$E[Y^{\mathrm{b}}_{i\in\mathrm{B}}|FTFY^{\mathrm{b}}_{i\in\mathrm{B}} = 1, S_i = s] > E[Y^{\mathrm{b}}_{i\in\mathrm{B}}|FTFY^{\mathrm{w}}_{i\in\mathrm{B}} = 1, S_i = s] \quad \text{(A8)}$$

and, even more abstractly, that:

$$E[Y^{\mathrm{w}}_{i\in\mathrm{B}}|FTFY^{\mathrm{b}}_{i\in\mathrm{B}} = 1, S_i = s] > E[Y^{\mathrm{w}}_{i\in\mathrm{B}}|FTFY^{\mathrm{w}}_{i\in\mathrm{B}} = 1, S_i = s]. \quad \text{(A9)}$$

For example, for class VIIa, if whites are more likely to be employed FTFY (e.g. by assumption A3) because blacks have higher reservation wages, then the average potential wages of blacks in class VIIa would be lower if blacks were employed as frequently as whites. Thus, the asymptotic bias of a conditional-on-S variant of estimator E1 could be either positive or negative. This is the claim of the contrarian selection narratives outlined in the main text, where the bias is declining in time, possibly generating illusory gains for blacks.

Table A1. Distribution of the Population of White and Black Males Across Types of Labor Market Participation.

	White Males (Age: 18–64)				Black Males (Ages: 18–64)			
	Proportion of CPS Sample		Mean Years Education		Proportion of CPS Sample		Mean Years Education	
	1985–1989	1996–2000	1985–1989	1996–2000	1985–1989	1996–2000	1985–1989	1996–2000
Worked one or more weeks last year								
Full-time, full-year	0.661	0.684	13.36	13.42	0.497	0.528	12.40	12.87
Full-time, part-year	0.159	0.122	12.41	12.56	0.175	0.124	11.82	12.40
Part-time, full-year	0.033	0.038	12.98	13.00	0.029	0.033	11.85	12.53
Part-time, part-year	0.046	0.039	12.68	12.79	0.065	0.044	11.50	12.18
Other (full-time and less than $50/week earnings)	0.004	0.005	12.47	12.37	0.003	0.004	11.33	12.20
Worked zero weeks last year								
Could not find work	0.008	0.006	11.03	11.60	0.036	0.020	11.11	11.62
Disabled	0.035	0.043	9.89	10.99	0.083	0.098	9.12	10.86
Keeping house	0.002	0.005	11.50	11.69	0.004	0.012	10.95	11.44
Retired	0.032	0.030	11.82	12.72	0.024	0.025	9.64	11.69
In school full-year	0.011	0.015	13.63	13.09	0.029	0.031	12.63	12.46
Other	0.004	0.006	11.44	11.58	0.014	0.017	10.63	11.52
Incarcerated	0.006	0.010			0.042	0.069		

To understand how important these complications may be, we first present basic results in Table A1 on race differences in labor market participation, again separately for the final five years of the two most recent economic expansions. The first 11 rows of the table are based on analysis of the entire CPS sample of black and white males between the ages of 18 and 64. For these groups of respondents, we also present in Table A1 measured levels of skill, as the mean years of education attained by respondents. The incarceration rates in row 12 of the table are drawn from the *Sourcebook of Criminal Justice Statistics 2000* (see http://www.albany.edu/sourcebook/1995/pdf/t619.pdf). The distribution across the first 11 rows within year-specific variants of Table A1 was determined by dividing the proportion of the CPS sample observed in each of the categories by $1/(1 +$ the race-specific incarceration rate$)$ in that year.[21]

As shown in the first row of Table A1, a greater proportion of the non-institutionalized population of both white and black males was employed FTFY in 1996–2000 than in 1985–1989. In particular, the proportion of white males working FTFY increased from 66.1 to 68.4% while the proportion of blacks working FTFY increased from 49.7 to 52.8% (which is a net gain of 0.7% for blacks). And, consistent with the greater relative aggregate upward mobility of black males over this time period demonstrated in Table 2, the measured skills of blacks increased more than those of whites, from 12.40 to 12.87 years of education for blacks and from 13.36 to only 13.42 years of education for whites.

For whites, the major difference across the two time periods appears to be movement out of part-year employment into full-year employment. FTPY employment decreased from 15.9 to 12.2% of all white males, and PTPY employment decreased from 4.6 to 3.9%. These trends can be attributed to the record low unemployment levels achieved in the late 1990s economic expansion. And for this reason as well, the core long-term unemployed, those who worked no weeks in the prior year but looked for work, decreased from 0.8 to 0.6%. For blacks, the same trends are evident, although the proportion of respondents in non-FTFY categories was generally higher than for whites in each time period. Nonetheless, FTPY employment decreased from 17.5 to 12.4% of all black males, and PTPY employment decreased from 6.5 to 4.4%. Likewise, the core long-term unemployed decreased from 3.6 to 2.0%. Even though blacks generally made gains relative to whites, this is somewhat misleading, as it is in part generated by the growing racial disparity in imprisonment. The incarceration rate for blacks increased from 4.2 to 6.9% while the incarceration rate for whites increased from 0.6 to 1.0%.

Can these trends account for a substantial portion of the relative decline in the race gap in classes V, VI, and VIIa? The type of estimator we deploy to address this possibility is specified in three main steps, all with the goal of generating a lower average estimate of the wages of blacks in classes V, VI, and VIIa in proportion to the lower propensity for blacks to be observed working as FTFY workers. First, all FTPY blacks were added to the analysis sample. Then, we calculated a set of scaling factors for black FTFY and FTPY workers within each class in each year. These scaling factors are functions of the observed rate of FTFY employment for whites within each year and a more complex "what if" rate of FTFY employment for blacks within each year, as specified below. Finally, we re-estimated all of the models in Table 3 with the sample that includes the FTPY black workers, weighted as specified below.

For comparison with Estimator E2, the selection-adjusted estimator can be written in general form as class and year-specific variants of:

$$E_N[\phi^{\mathrm{w}}(k,t)Y_i|B_i=0, FTFY_i=1] \\ - E_N[\phi^{\mathrm{b}}(k,t)\psi^{\mathrm{b}}(c,t)Y_i|B_i=1, FTFY_i=1 \text{ or } FTPY_i=1]. \qquad \text{(E3)}$$

Notice that the first term of estimator E3 is nearly identical to the first term of estimator E2, and only differs because the term $\phi^{\mathrm{w}}(k,t)$ is utilized instead of $\pi^{\mathrm{w}}(k,t)$.[22] The selection adjustment is generated by the alternative "what if" estimate for blacks in the second term. The key to the selection adjustment is the scaling factor $\psi^{\mathrm{b}}(c,t)$ for blacks, which, as defined in the main text, is a function of $\lambda^{\mathrm{w}}(c,t)$, $\lambda^{\mathrm{b}}(c,t)$, and $\theta^{\mathrm{b}}(c,t)$. Averaged across the two time periods within each class, the values of these component rates are presented in Table A2, along with the weekly earnings of FTFY and FTPY blacks. The rates $\lambda^{\mathrm{w}}(c,t)$ and $\lambda^{\mathrm{b}}(c,i)$ are straightforward, as they are simply the year and class-specific percentages of observed FTFY and FTPY workers who are observed FTFY workers, respectively for white and black men. As shown in the first four columns of Table A2, values for these rates are mostly between 0.7 and 0.9 and are highest for classes I, II, and V. Likewise, $\lambda^{\mathrm{w}}(c,t)$ is more often than not greater than $\lambda^{\mathrm{b}}(c,t)$, in part because blacks are more likely to experience unemployment spells.

The key to the selection adjustment, however, is the "what if" rate of FTFY employment, $\theta^{\mathrm{b}}(c,t)$. This rate is calculated for a wider population of blacks, augmented proportionally to the greater representation of whites in FTFY employment. This "what if" rate is generated in three steps:

(1) Calculate the gross additional percentage of blacks who would be FTFY workers if the black rate of FTFY employment was as high as the white rate of FTFY employment. For the categories in Table A1, there are six distinct groups in each year t, defined as the proportion of blacks minus the proportion of whites who: (1) Worked zero weeks last year because they could not find work; (2) worked zero weeks last year because they were disabled; (3) worked zero weeks last year because they were keeping house; (4) worked zero weeks

Table A2. Breakdown of Specific Components of Estimator E3 by Class and Time Period.

	$\lambda^{w}(c)$		$\lambda^{b}(c)$		$\theta^{b}(c)$		Log Weekly Earnings of Black, FTFY Workers		Log Weekly Earnings of Black, FTPY Workers	
	1985–1989	1996–2000	1985–1989	1996–2000	1985–1989	1996–2000	1985–1989	1996–2000	1985–1989	1996–2000
I	0.914	0.926	0.913	0.918	0.897	0.891	6.80	6.89	6.54	6.58
II	0.889	0.905	0.892	0.884	0.878	0.859	6.43	6.51	6.32	6.38
IIIa	0.834	0.851	0.817	0.823	0.806	0.801	6.27	6.36	5.94	5.93
IIIb	0.741	0.780	0.651	0.775	0.637	0.753	5.97	6.06	5.56	5.80
V	0.883	0.897	0.876	0.894	0.859	0.861	6.42	6.54	6.19	6.35
VI	0.743	0.814	0.715	0.811	0.698	0.769	6.27	6.32	6.02	6.16
VIIa	0.698	0.775	0.708	0.774	0.681	0.709	6.05	6.08	5.77	5.85

last year for “other reasons”; (5) worked zero weeks last year because they were incarcerated; and (6) worked full-time last year for some amount of weeks but earned on average less than 50 dollars per week.

(2) Assume that these additional blacks would be employed according to the following class distribution: 0.05 for class I, 0.08 for class II, 0.04 for class IIIa, 0.015 for class IIIb, 0.045 for class V, 0.12 for class VI, and 0.65 for class VIIa. This distribution is arbitrary, but it is similar to the distribution of black employment in Table 2, although weighted more heavily toward class VIIa. Then, multiply the two distributions together: the year-specific percentage of gross additional blacks and the stipulated “what-if” class distribution of employment for the gross additional blacks. The result is a distribution of additional blacks, denoted as a set of proportions $\eta^{b}(c, t)$ which vary over classes c and years t.

(3) Define an adjusted rate of observed FTFY employment among a more broad-based set of blacks, assuming that there are now additional blacks in the denominator:

$$\theta^{\mathrm{b}}(c, t) = \frac{\omega^{\mathrm{b}}(c, t)\lambda^{\mathrm{b}}(c, t)}{\omega^{\mathrm{b}}(c, t) + \eta^{\mathrm{b}}(c, t)}$$

where $\omega^{\mathrm{b}}(c, t)$ is the observed percentage of all black FTFY and FTPY CPS respondents who are in class c in year t. In all years, $\theta^{\mathrm{b}}(c, t)$ is lower than $\lambda^{\mathrm{b}}(c, t)$ for all classes.

Finally, the three rates $\lambda^{\mathrm{w}}(c, t)$, $\lambda^{\mathrm{b}}(c, t)$, and $\theta^{\mathrm{b}}(c, t)$ are then combined into the selection-adjustment scaling factor $\psi^{\mathrm{b}}(c, t)$ that is applied to black FTFY and FTPY workers in the augmented sample. As the difference between $\lambda^{\mathrm{w}}(c, t)$ and $\theta^{\mathrm{b}}(c, t)$ grows, the amount of weight given to the lower FTPY wages of blacks increases, in line with the selection narratives outlined in the main text.

A key assumption of the selection-adjustment framework is that, within class c and year t, the additional “what if” blacks summarized in the aggregate by $\eta^{\mathrm{b}}(c, t)$, would on average receive lower wages than those blacks who are observed working FTFY. In particular, we assume that they would receive the wages typical of black FTPY workers, which are on average 15–20% lower (see Table A2). Since there is virtually no way to know what these hypothetical “what if” black workers would earn if they were employed FTFY, we have no firm justification for this assumption. We regard it as a rather pessimistic assumption, since one might instead assume that if these “what if” black workers were actually employed FTFY they might be earning wages much closer to what observed black FTFY workers typically earn (which would still place their wages below those of whites with comparable measured covariates). Despite our judgement, we experimented with even more pessimistic assumptions. In these ad hoc explorations, our models remained surprisingly robust, primarily because the trends in the black-white gap are largely a function of the consistent relative decline in the earnings of white men in classes V, VI, and VIIa.

NOTES TO APPENDIX

20. Again, $E_N[\cdot]$ denotes the sample expectation. Thus, assumption A1, for example, states that, as N increases to infinity, an implied sequence of sample estimates, compactly denoted as $E_N[Y_i|B_i = 0]$, converges in probability to a precisely defined conditional population expectation $E[Y^{\mathrm{w}}_{i \notin \mathrm{B}}]$ that is regarded as a theoretical constant.

21. The incarceration rate is the percentage of the adult population, respectively for whites and blacks, resident in federal and state prisons and local jails. We used raw values from 1985 to 1997, and then formed linear extrapolations for 1982–1984 and 1998–2000 from 1985–1988 to 1994–1997, respectively.

22. The weights $\phi^{\mathrm{w}}(k, t)$ and $\phi^{\mathrm{b}}(k, t)$ are nearly identical to $\pi^{\mathrm{w}}(k, t)$ and $\pi^{\mathrm{b}}(k, t)$, except that they calculate the percentage representation in each occupation using both black FTFY and FTPY workers. When not invoking the non-parametric occupational alignment, they are implicitly set equal to 1, as in models in the first column of each panel of Table 4. For other models, they take on values in order to align the occupational distributions of whites and blacks to the occupational distribution of blacks in the chosen reference time period, as in the models in the second and third column of each panel in Table 4.