WHEN WORK MOVES: JOB SUBURBANIZATION AND BLACK EMPLOYMENT

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Abstract—This paper examines whether job suburbanization caused declines in black employment rates from 1970 to 2000. I find that black workers are less likely than white workers to work in observably similar jobs that are located further from the central city. Using evidence from establishment relocations, I find that this relationship reflects at least in part the causal effect of job location. At the local labor market level, I find that job suburbanization is associated with substantial declines in black employment rates relative to white employment rates. Evidence from nationally planned highway infrastructure corroborates a causal interpretation.

I. Introduction

OVER the past 60 years in the United States, the unemployment rate among black working-age adults has been roughly double the national unemployment rate (Fairlie & Sundstrom, 1999). The *spatial mismatch hypothesis*, introduced by Kain (1968) and further popularized by Wilson (1987, 1996), attributes racial disparities in employment rates in part to spatial frictions in the labor and housing markets. Under this theory, black households tend to live relatively far from work opportunities, reducing their access to gainful employment. This distance increased after World War II, as firms and white households began relocating from central cities to suburban rings at an accelerated pace.[1] Black households, who faced discrimination in housing and mortgage markets (Yinger, 1995; Rothstein, 2017), remained concentrated in central cities.[2] As a result, black households tend to live further away from the portions of metropolitan areas experiencing substantial job growth, depressing their labor market outcomes. For example, between 1970 and 2000 in the sample of metropolitan areas studied here, the employment rate for black men declined by about 10% more than the employment rate for white men.[3]

It remains unclear whether job suburbanization and the decline of black employment over this period are causally linked, however. Though researchers have posited several explanations for why work has decentralized, including improvements in transportation technology and infrastructure, land costs, and worker suburbanization (Glaeser & Kahn, 2001), the process of job suburbanization is not well understood. Job suburbanization may be associated with changes in labor demand and supply that would generate changes in the racial composition of the workforce even in the absence of suburbanization. Moreover, even if firms began to locate in the suburbs due to exogenous factors, it is not clear what implications this would have for the racial composition of the workforce. While the segregation of black households in the central city has been a persistent feature of U.S. metropolitan areas (Cutler et al., 1999), workers may respond to the changing geography of work by changing jobs, altering their commute, or moving to a different neighborhood. These response margins may be sufficient for job suburbanization to have little effect on labor market outcomes by race.

In this paper, I examine whether job suburbanization caused significant declines in black employment from 1970 to 2000. I provide both job-level and local labor market-level evidence. At the job level, I find that black workers are less likely than white workers to work in observably similar jobs located further from the central city. Using evidence from establishment relocations, I find that this relationship reflects at least in part the causal effect of job location. At the local labor market level, I find that job suburbanization is associated with substantial declines in black employment rates relative to white employment rates. Evidence from nationally planned highway infrastructure corroborates a causal interpretation. My findings imply that job suburbanization can explain the majority of the relative decline in black men's employment over this period.

After I introduce the data and how I measure job suburbanization (section II), the analysis is divided into two sections. In the job-level analysis (section III), I use establishment-level administrative data from the Equal Employment Opportunity Commission to show that black workers are substantially less likely than white workers to work in jobs located further from the central city. In particular, conditional on job characteristics, the black share of employees in metropolitan area establishments is decreasing in an establishment's distance from the central business district (CBD). Remarkably, this spatial segregation is stable over time despite widespread movement of population and jobs to the suburbs. I also provide additional evidence that this spatial segregation reflects at least in part the *causal* effect of job location on racial composition. I show that the relationship between job location and racial composition holds *within firms* that operate multiple establishments within a metropolitan area. I also find that establishments that relocate from the central city to the suburbs experience sharp coincident declines in the black share

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[1] In 1960, 61% of metropolitan area jobs were in the central city; by 2000, the central city share declined to 34% (Baum-Snow, 2020).

[2] See the online appendix for additional details.

[3] Among women, the employment rate for white women increased by about 34% more than the employment rate for black women.

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of their employees despite no significant changes in their occupational composition.

This persistent spatial segregation suggests job suburbanization may have reduced black employment rates. Effect size estimates from establishment relocations suggest that job suburbanization decreased the black share of the workforce by 2.3% each decade. However, suburbanization may be offset in the aggregate by worker resorting across workplaces. This motivates the second portion of the analysis (section IV), which examines the relationship between job suburbanization and black employment across local labor markets.

Using census data and a synthetic panel, differences-in-differences research design, I find that job suburbanization is associated with substantial declines in black employment rates relative to white employment rates. For every 10% decline in the fraction of metropolitan statistical area (MSA) jobs located in the central city over this period, black relative employment rates declined by 1.6%–2.3%, while white employment rates increased by a (statistically insignificant) 0.3%–0.4%. This relationship holds within observable skill groups, and estimates are similar for men and women. Relative earnings declined by 1.2%–2.3%, though these estimates are less stable across specifications.

To address the potential endogeneity of job suburbanization—in particular, changes in the spatial distribution of work driven by unobserved labor supply shocks that are unevenly distributed across black and white working-age adults—I instrument for job suburbanization using variation in nationally planned interstate highway construction across MSAs as identified in Baum-Snow (2007). These highways were planned in the 1940s and 1950s and were primarily designed to link faraway places rather than to facilitate local commuting or economic development. Hence, their assignment across MSAs should be exogenous to residual labor supply shocks from 1970 to 2000. While highways have a variety of effects on the labor market that may potentially violate the exclusion restriction, I argue that they are unrelated to labor demand and supply shocks that would disproportionately affect black workers, the most concerning omitted variables. In particular, suburbanization induced by highway construction is not related to changes in local industry or occupation mix that would portend changes in black relative employment rates.

Consistent with Baum-Snow (2020), I find that each highway ray emanating from the central city leads to a 7%–10% decline in the fraction of MSA jobs located in the central city from 1970 to 2000. In turn, interstate highways caused declines in black relative employment rates, with highway-based instrumental variable (IV) estimates for the relationship between job suburbanization and black employment rates that are comparable to the ordinary least squares (OLS) estimates. I conclude that job suburbanization was an important determinant of black-white labor market inequality from 1970 to 2000. The estimates imply that job suburbanization can explain over half of the relative decline in black men's employment rates and 15%–20% of the increase in white women's employment rates relative to black women's over this period.

While this paper provides novel evidence on spatial segregation in the labor market and the causal effect of workplace location on the racial composition of employees, it is largely silent on why space matters. I discuss potential mechanisms in section V. A key implication of this paper is that the relationship between job suburbanization and black employment is not driven by changes in firm demand for worker skills. Interestingly, the labor market is substantially less spatially segregated than the housing market, so commuting flows do offset residential segregation to some degree. However, while black households are residentially less concentrated in central city neighborhoods than they were in 1970, this movement has not been sufficient to noticeably alter the spatial segregation of the labor market.

This paper contributes to an extensive literature testing the spatial mismatch hypothesis. Prior work typically relates labor market outcomes to measures of job accessibility in a cross-section (see Ihlanfeldt and Sjoquist, 1998, for a review of the older literature).[4] Most recent work in this literature finds some support for spatial mismatch, though there has been considerable debate about its empirical importance (Ellwood, 1986). The results tend to be sensitive to how job accessibility is measured (Raphael, 1998).[5] More importantly, results from this literature are made difficult to interpret by the endogeneity of household and firm location. Residents who are less productive may sort into neighborhoods farther from work opportunities, where rents are typically lower.[6] Similarly, firms may choose to locate in neighborhoods with residents who are more productive. In this paper, I take a more "reduced-form" approach—rather than attempt to estimate the effects of work proximity per se, I estimate the effects of job suburbanization at the local labor market level.[7]

This paper also contributes to a smaller literature that studies how a work establishment's location influences the racial composition of its employees. This research has found that location appears to be an important determinant of employee composition. However, prior work has been limited to

[4]Three exceptions are Mouw (2000), Weinberg et al. (2004), and Stoll (2006). Mouw (2000) estimates the relationship between changes in job density and neighborhood-level employment rates in Chicago and Detroit from 1980 to 1990. Weinberg et al. (2004) exploit individual moves across neighborhoods using the 1979 National Longitudinal Survey of Youth. Both papers find evidence of spatial mismatch. Stoll (2006) relates changes in job sprawl to changes in spatial mismatch between where black households reside and employers are located across MSAs from 1990 to 2000 and finds no detectable relationship.

[5]For example, previous researchers have used the local job density, local job growth, and the average commuting times of local workers as measures of job accessibility.

[6]Alternatively, if spatial frictions are relevant, residents who find it difficult to obtain work may sort into neighborhoods with higher employment density.

[7]Andersson et al. (2018) is an important and recent contribution to this literature. The authors use matched employer-employee data to study the relationship between job accessibility and jobless duration among workers displaced in mass layoffs. They find that, among similar job searchers, search duration is decreasing in job accessibility.

relatively small samples—cross-sectional studies of a few thousand firms in a handful of metropolitan areas (Holzer & Ihlanfeldt, 1996) or case studies of individual plant relocations (Zax & Kain, 1996; Fernandez, 2008). By contrast, the administrative data used here cover hundreds of thousands of establishments for several decades, including thousands of relocation episodes.

II. Data and Measurement

In this paper, I use three data sources: establishment-level data from EEO-1 forms, individual-level and city-level census data, and MSA-level data on the interstate highway system from Baum-Snow (2007). In this section, I describe each data source and introduce my MSA-level measure of job suburbanization.

A. *EEO-1 Form Data*

For the job-level analysis, I use a unique set of establishment-level panel data. These data, known as EEO-1 form data, are collected by the U.S. Equal Employment Opportunity Commission (EEOC) and cover the years 1971–2000. As part of the Civil Rights Act of 1964, firms meeting certain size requirements are required to complete EEO-1 forms annually and submit them to the EEOC. Firms are required to report their overall racial, ethnic, and gender composition and the racial, ethnic, and gender composition of each of their establishments meeting size requirements, disaggregated by nine major occupation groups.[8] Employers are instructed to base demographic classifications on worker self-identification or visual inspection, where the former is the preferred method.[9]

Before 1982, all firms with 50 or more employees were required to submit EEO-1 forms. In 1982, the firm size cutoff was adjusted up to 100. For the analysis, I drop all establishments that would not meet the post-1982 criteria. Firms are required to file a separate report for each establishment with at least 50 employees and the company headquarters. Establishments are consistently identified with firm and establishment identifiers, and I observe each establishment's location and industry. Online appendix table A3 presents summary statistics for the EEO-1 data covering the 58 MSAs studied here. For most of the analysis using EEO-1 data, I restrict to establishments that I can geocode using street address, zip code, or city.[10] The results are very similar if I restrict to establishments that I can geocode using street address.[11]

B. *Census Data*

For the local labor market-level analysis, I use data from the decennial census. I use census data from 1970, 1980, 1990, and 2000 to measure various labor market characteristics and job suburbanization. I focus on these census years for two reasons. First, the second wave of the Great Migration, a period when a substantial share of Southern black households moved to other regions of the country, ends around 1970. Analysis of census data from earlier than 1970 would be complicated by the large and potentially endogenous black migration flows over this period. Second, MSA is not identified in the publicly available 1960 census microdata.

To measure labor market characteristics, I use Integrated Public Use Microdata Series (IPUMS) census data (Ruggles et al., 2010). The 1970 census data are a 2% national sample, while the remaining years are 5% national samples. I restrict the analysis to the 58 consistently identified MSAs with the largest black populations as defined in 1970.[12] These MSAs are listed in online appendix table A1.

To measure job suburbanization, I use various census data products. Measuring the spatial distribution of work consistently across years is complicated by the fact that central city definitions change significantly with the 1990 census. In particular, many cities defined as suburbs in 1980 are classified as central cities in 1990. These changes make it difficult to construct consistent measures of job suburbanization after 1980 using only IPUMS census data. Instead, I combine IPUMS data with tabulations from the Census Transportation Planning Package (CTPP) to measure the spatial distribution of work in 1990 and 2000.[13] The CTPP data include tabulations reporting the total number of individuals working at various levels of geography. I divide those totals into central cities and suburbs using 1970 census definitions for central cities. For 1970 and 1980, I use the IPUMS census data. In the census microdata, I measure the spatial distribution of work using the census indicator for whether an individual works in the central city or outside the central city (suburbs) of an MSA. Note that while I hold the set of municipalities defined as central cities constant, I follow census definitions of central city and MSA geographies, which evolve over time in some cases.

[8]The nine occupation categories consist of the following: officials and managers, professionals, technicians, sales workers, administrative support workers, craft workers, operatives, laborers/helpers, and service workers.

[9]There is no distinction between race and ethnicity in the data; in particular, Hispanic workers are classified as a distinct, nonoverlapping group.

[10]For establishments that I can only geocode using zip code or city, I assign the coordinates of the centroid for that zip code or city.

[11]Due to the size requirements, establishments in the EEO-1 data are not representative of all U.S. establishments. Industries that tend to have large establishments (e.g., manufacturing) are overrepresented, while industries that tend to have small establishments (e.g., services) are underrepresented. Overall, the EEO-1 data account for about 40% of total employment (Robinson et al., 2005). Black workers are overrepresented at large establishments and firms, but this overrepresentation appears to be stable over this period (Carrington et al., 2000).

[12]Unfortunately, after 1970, many MSAs are only partially identified in the IPUMS census data. That is, some MSA residents are not identified as MSA residents in the data. These residents tend to live in (suburban) areas that straddle MSA boundaries, so the representativeness of the black population (who tend to reside in central cities) should be less affected. Nevertheless, I restrict to MSAs where no more than 15% of the estimated MSA population is omitted in 1980, 1990, or 2000.

[13]This follows Baum-Snow (2020), who uses the CTPP from 2000 to measure commuting flows between suburbs and central cities.

C. Interstate Highway Data

I use data on the number of interstate highway rays emanating from MSA central cities in 1970 and the radius of the central city as measured in 1950. These data are collected in Baum-Snow (2007). I use data on MSA exposure to the interstate highway system as a source of variation in job suburbanization.

D. Measuring Job Suburbanization

To analyze the effects of job suburbanization using variation across MSAs, I need a measure of job suburbanization that can be applied consistently across MSAs with differing initial spatial distributions of work. I also need a measure that can be calculated using available census data. I focus on the proportional change in the number of central city jobs due to the change in the spatial distribution of work, what I term the *share effect*. The number of central city jobs may change because the whole MSA is growing or shrinking—the *scale* effect—or via the share effect. More concretely, let T_t denote the number of jobs in an MSA at time t ($t = 70, 80, 90, 00$), let π_t^{cc} denote the fraction of MSA jobs located in the central city, and let T_t^{cc} denote the number of central city jobs. The change in the log of the number of central city jobs can be decomposed as follows:

$$\begin{aligned}\log T_{t_1}^{cc} - \log T_{t_0}^{cc} &= \log(\pi_{t_1}^{cc} T_{t_1}) - \log(\pi_{t_0}^{cc} T_{t_0}) \\ &= \underbrace{[\log \pi_{t_1}^{cc} - \log \pi_{t_0}^{cc}]}_{\textit{share}\text{ effect}} + \underbrace{[\log T_{t_1} - \log T_{t_0}]}_{\textit{scale}\text{ effect}}.\end{aligned}$$

Hence, I measure job suburbanization using $\Delta_{t_1,t_0} \log \pi^{cc} = \log \pi_{t_1}^{cc} - \log \pi_{t_0}^{cc}$. In each decade, the fraction of work in the central city (π^{cc}) decreases by an average of about 10% across MSAs. However, this average masks substantial heterogeneity across MSAs; the standard deviation is about 10% over each decade. I report job suburbanization for all MSAs included in the analysis in online appendix tables A1 and A2.

III. Job Location and Racial Composition

In this section, I assess whether, conditional on job characteristics, black workers are less likely to work in the suburbs than white workers and how the relationship between job location and racial composition has changed over time. I use establishment relocations to isolate the causal effect of job location on racial composition.

I use EEO-1 data to measure how the racial composition of workplaces varies with workplace location. The granularity of the EEO-1 data allows for a finer measure of location than an indicator for central city status. Ideally, I would measure location effects on racial composition for very disaggregated areas within metropolitan areas (e.g., census tracts), adjusting for establishment characteristics. Unfortunately, there are not enough establishments in the EEO-1 data to do this effectively. Instead, I parameterize the effect of location, focusing on an establishment's distance from the corresponding CBD of the central city.[14]

To facilitate comparisons across MSAs, I first normalize each establishment's black share of employees by the black share of employees for all establishments in the MSA that year. I refer to this measure as an establishment's *normalized black share*. Hence, in an MSA where 15% of the workforce is black, an increase of 0.10 in an establishment's normalized black share maps to a 1.5 percentage point increase in the establishment's black share of employees.

Panel A of figure 1 plots the relationship between an establishment's distance from the CBD and normalized black share from 1971 to 1975 and from 1996 to 2000. I plot local linear fits and confidence bands, where establishments are weighted by their number of employees.[15] In both periods, there is a distinct negative relationship between distance from the CBD and normalized black share. For every 10-mile increase in the distance from the CBD, an establishment's normalized black share decreases by 0.25. Remarkably, this relationship has changed little from 1971 to 2000. By contrast, the analogous slope for the racial makeup of residential neighborhoods (as measured by census tracts), depicted in panel B of figure 1, is substantially steeper but flattens significantly over this period.[16] Hence, while worker commuting patterns generally mute the translation of residential segregation into workplace segregation, the spatial segregation of the labor market remained roughly constant over this period despite significant residential desegregation.

The negative relationship between an establishment's distance from the CBD and its black share of employees suggests that, within an MSA, job location is an important determinant of racial composition. However, it may also reflect the fact that location is correlated with other important determinants of racial composition, including industry, occupation, and establishment size. To adjust for these job characteristics, I estimate the following model:

$$\begin{aligned}&\text{norm. black share}_{jt} \\ &= \tau_{m(j)i(j)t} + \beta \text{distance}_{jt}^{CBD} + \gamma \log(\text{est. size})_{jt} + \epsilon_{jt}, \quad (1)\end{aligned}$$

where j indexes establishments, $i(\cdot)$ indexes industries, $m(\cdot)$ indexes MSAs, and $\tau_{m(j),i(j),t}$ are MSA-by-industry-by-year fixed effects. In one specification I replace the MSA-by-industry-by-year fixed effects with MSA-by-*firm* by year fixed effects. Under this specification, β is identified from

[14]The Census Bureau defines a CBD as "an area of very high land valuation characterized by a high concentration of retail businesses, service businesses, offices, theaters, and hotels, and by a very high traffic flow." I use the latitude and longitude of each CBD as measured in Holian and Kahn (2012).

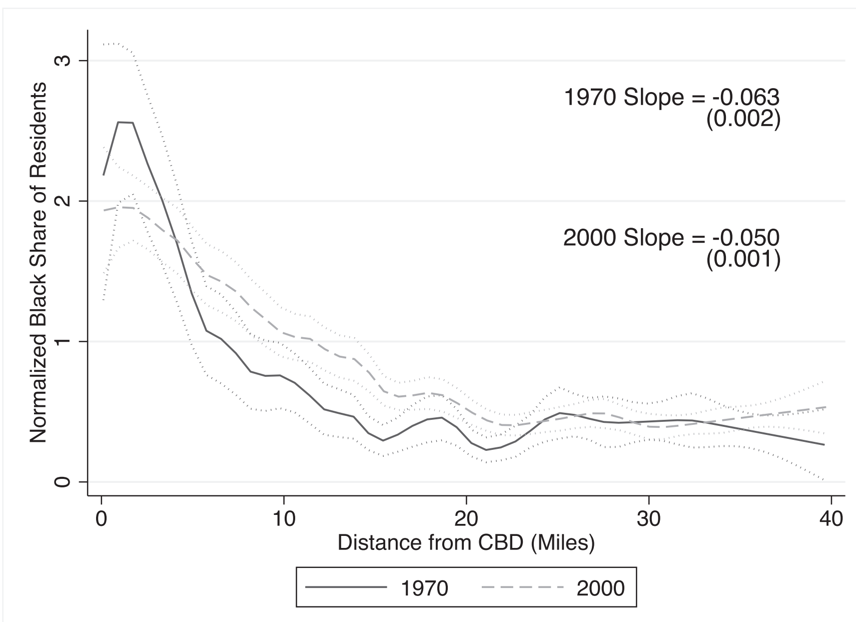
[15]I use the Stata package `binsreg` developed by Cattaneo et al. (2019) to compute the local linear fits and confidence bands.

[16]Census tracts are weighted by population. This pattern is consistent with Cutler et al. (1999), who document significant declines in residential segregation from 1970 to 2000 as measured by the dissimilarity and isolation indices.

FIGURE 1.—DISTANCE FROM CBD AND BLACK SHARE OF EMPLOYEES AND RESIDENTS

(A) Establishments

(B) Neighborhoods



Panel A plots the nonparametric relationship between an establishment's normalized black share of employees and its distance from the central business district, weighting by establishment size. Panel B plots nonparametrically the relationship between a neighborhood's (as measured by census tracts) normalized black share of residents and its distance from the central business district, weighting by tract population. I plot local linear fits and confidence bands using the Stata package `binsreg` developed by Cattaneo et al. (2019).

variation in establishment location in the same MSA and year *within the same firm.* I weight observations by establishment size. I also estimate analogous models where the observations are at the *job* level (establishment-by-occupation), where I include MSA-by-industry-by-occupation-by-year fixed effects (or firm-by-occupation-by-year fixed effects) and weight observations by job cell size. In all models, I cluster standard errors at the establishment level.

Table 1 presents slope estimates. Panel A presents establishment-level estimates, and panel B presents job-level estimates. In column 1, I pool all years of data and include MSA-by-industry-by-year fixed effects. The estimated coefficient, −0.0251, implies that for every 10-mile increase in its distance from the CBD, an establishment's normalized black share decreases by 0.251. To put this magnitude in perspective, note that in the EEO-1 data, the average distance from the

TABLE 1.—DISTANCE FROM CBD AND BLACK SHARE OF EMPLOYEES

Outcome: Normalized Black Share Panel A: Establishment	All		By Decade		
			1970s	1980s	1990s
	(1)	(2)	(3)	(4)	(5)
Distance from CBD (Miles)	−0.0251**	−0.0283**	−0.0250**	−0.0252**	−0.0250**
	(0.0004)	(0.0007)	(0.0006)	(0.0005)	(0.0004)
log Establishment Size	✓	✓	✓	✓	✓
Industry × MSA × Year FEs	✓		✓	✓	✓
Firm × MSA × Year FEs		✓			
# Establishments	418,906	418,906	193,401	187,369	210,097
Panel B: Within-Occupation	All		By Decade		
			1970s	1980s	1990s
	(6)	(7)	(8)	(9)	(10)
Distance from CBD (Miles)	−0.0271**	−0.0324**	−0.0274**	−0.0273**	−0.0267**
	(0.0004)	(0.0007)	(0.0005)	(0.0004)	(0.0004)
log Establishment Size	✓	✓	✓	✓	✓
Ind. × Occ. × MSA × Year FEs	✓		✓	✓	✓
Firm × Occ. × MSA × Year FEs		✓			
# Establishments	418,906	418,906	193,401	187,369	210,097

This table refers to estimates of equation (1). All models include log establishment size as a control. Panel A models are estimated at the establishment level. Panel B models are estimated at the job level (establishment by occupation). Standard errors are in parentheses, clustered at the establishment level. Regression is weighted by establishment size (panel A) or job cell size (panel B). ~Denotes statistical significance at the $p < 0.10$ level. *Denotes statistical significance at the $p < 0.05$ level. **Denotes statistical significance at the $p < 0.01$ level.

CBD is 6 miles for central city establishments and 18 miles for suburban establishments, a difference of 12 miles. Column 2 includes MSA-by-firm-by-year fixed effects, and the coefficient *increases* in magnitude to −0.0283. In columns 3–5, I estimate equation (1) by decade, including MSA-by-industry-by-year fixed effects. The estimates are stable across time periods. The analogous within-occupation estimates in panel B are similar, though they are slightly larger in magnitude. Controlling for fine job characteristics, black workers are less likely to work in the suburbs than white workers, and this relationship between location and racial composition is stable over time.

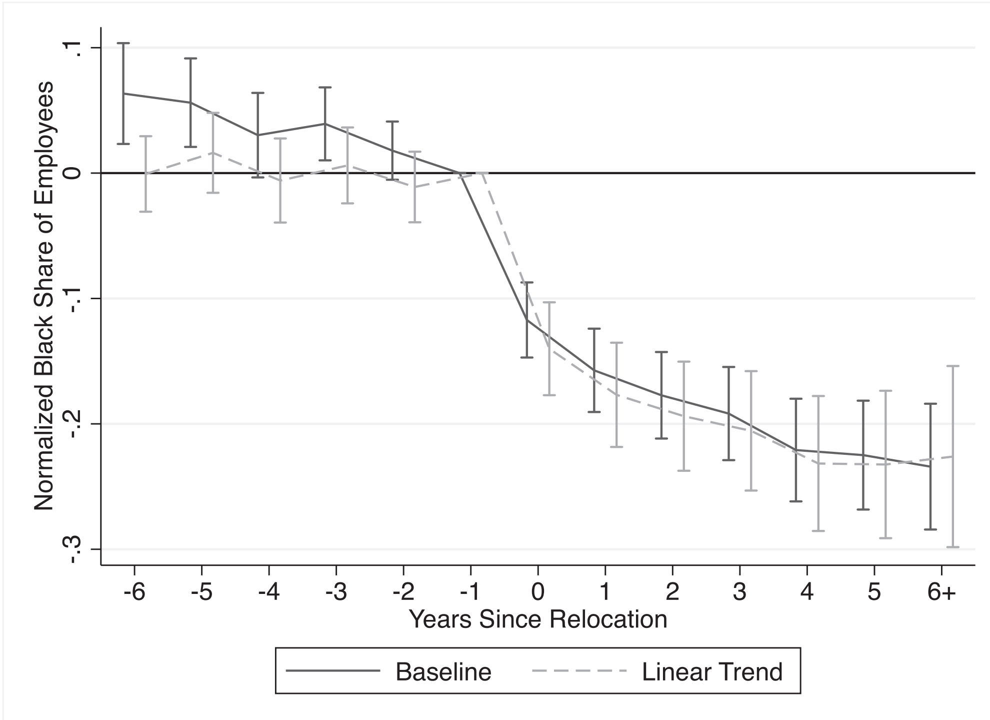
The cross-sectional relationship between distance from the CBD and normalized black share suggests that spatial frictions play a significant role in determining where people work. However, it may be the case that jobs located further from the CBD differ in important unobserved ways so that, independent of location, those jobs would be less likely to be filled by black workers. These may include characteristics of the work itself or establishment-level preferences over workers. To provide additional evidence that spatial frictions play a significant role in the racial composition of an establishment's workforce, I estimate the effect of an establishment's relocation on its black share of employees. The advantage of studying establishment relocations is that job characteristics and local labor market conditions are (approximately) fixed before and after the relocation. As long as relocations are not associated with other types of restructuring that alters the racial mix of employees—an assumption I revisit below—any change in the racial composition of employees following the move should be due to the establishment's *location*.

This empirical strategy follows the work of Zax and Kain (1996) and Fernandez (2008). Both papers are case studies of single manufacturing plants that relocate from central cities (Detroit in Zax and Kain, 1996; Milwaukee in Fernandez, 2008) to neighboring suburbs and study how plant employees respond to those relocations. In both papers, the authors use worker-level personnel data and estimate models for the decision to quit the job and the decision to move to a new address. My approach here differs from prior work in several important ways. First, I have data on about 1,500 establishment relocations spanning 1972 to 2000. With data on significantly more relocations, I can make more general statements about the effects of relocation. Moreover, while prior work has relied on before and after snapshots, I use yearly panel data, allowing for a more credible event study research design. Unfortunately, while Zax and Kain (1996) and Fernandez (2008) use worker-level data, I only have access to establishment-level data on workforce composition. Hence, I cannot measure how worker decisions depend on worker-specific changes in commuting time. Instead, I will look at how the composition of the *entire establishment* changes with relocation.

I restrict the analysis to establishments that (1) move from a central city to a suburb within a given MSA and whose distance from the central city's CBD increases by at least 5 miles, or (2) remain in the same central city. I identify 1,501 establishments meeting these criteria, with an average increase in distance from the CBD of 12.5 miles.[17] This is similar to the difference in average distance from the CBD between central city and suburban establishments. In online appendix table A3, I present descriptive statistics for the establishments in the estimation sample.

[17] By contrast, I identify only 44 establishments that relocate from a suburb to the central city where the distance from the CBD decreases by at least 5 miles.

FIGURE 2.—BLACK SHARE DROPS FOLLOWING ESTABLISHMENT RELOCATIONS TO SUBURBS



This figure plots event study coefficients (θ_k) and 95% confidence intervals (dotted) estimated using models (2) and (3) where the outcome variable is the establishment's normalized black share of employees. The models are estimated using 1) all establishments that relocate from the central city to the suburbs and whose distance from the central business district increases by at least five miles and 2) establishments that remain in the central city. The coefficient for the year prior to the event (θ_{-1}) is normalized to zero. Estimated models include census division by industry by year fixed effects and log establishment size as controls. Standard errors are clustered at the establishment level. The green line depicts estimates derived from equation (2). The orange line depicts estimates derived from equation (3), which includes a linear trend specific to relocating establishments.

I estimate event study regression models of the following form, using data from both relocating establishments and central city establishments that do not relocate from the central city:

$$(\text{norm. black share})_{jt} = \alpha_j + \lambda_{d(j),i(j),t} + \gamma \log(\text{est. size})_{jt} + \sum_{k=a}^{b} \theta_k D_{jt}^k + \epsilon_{jt}, \quad (2)$$

where α_j and $\lambda_{d(j),i(j),t}$ are establishment and census division by industry by year fixed effects. D_{jt}^k are leads and lags for the relocation of establishment j. Let τ_j denote the year that establishment j relocates. Then D_{jt}^k are defined as

$$D_{jt}^k = \begin{cases} \mathbf{1}(t \leq \tau_j + a) & \text{for } k = a, \\ \mathbf{1}(t = \tau_j + j) & \text{for } a < k < b, \\ \mathbf{1}(t \geq \tau_j + b) & \text{for } k = b. \end{cases}$$

I normalize the value of $\theta_{-1} = 0$. The sequence of θ_j can be interpreted as the difference in establishment black share from the year prior to relocation and j periods thereafter, relative to nonrelocating establishments. Note that in this baseline specification, the end points pool the end point years (a or b) and years further from relocation ($< a$ or $> b$). For estimation, I set $a = -6$ and $b = 6$. For nonrelocating establishments, all the D_{jt}^k are set to zero. The identifying assumption is that, if not for relocation, the normalized black shares of relocating and nonrelocating establishments would be on parallel trends. Below I assess the plausibility of this assumption by examining the relative trends of relocating establishments prior to relocation.

I plot the θ coefficients in figure 2, and the pattern is stark. Prior to relocation, establishments exhibit little evidence of pretrends, though their normalized black share drops by about 0.03 from three years prior to the move to one year prior. This slight drop may reflect a trend that would continue even in the absence of relocation if, for example, relocating employers are already shedding black employees. It may also reflect workers anticipating the future move. In the year of relocation, normalized black share drops sharply by 0.12. Hence, even if relocating employers would have reduced their number of black employees if they had not relocated, this pattern indicates that relocation alone *causes* a decline in the black share of employees. This decrease continues following the move so that six years after the move, the normalized black share has dropped by about 0.23. On average, movers had a normalized black share of 0.85 in the year prior to the move.

Given suggestive evidence that relocating and nonrelocating establishments are on differential trends, and this differential trend is approximately linear, I estimate an alternative specification that allows for a linear trend specific to relocating establishments. In particular, I estimate

$$
\begin{aligned}
&(\text{norm. black share})_{jt} \\
&= \alpha_j + \lambda_{d(j),i(j),t} + \delta(\text{ever relocate}_j) \\
&\quad \times t + \gamma \log(\text{est. size})_{jt} + \sum_{k=a}^{b} \theta_k D_{jt}^k + \epsilon_{jt}, \qquad (3)
\end{aligned}
$$

where $(\text{ever relocate})_j$ is an indicator for whether establishment j ever relocates from the central city and $(\text{ever relocate})_j \times t$ is a differential time trend for relocating establishments. In this specification I adjust the definition of D_{jt}^{-6}, the end-point lead for relocation, to $D_{jt}^{-6} = \mathbf{1}(t = \tau_j - 6)$. In words, I no longer pool 6 years prior to the relocation with earlier years so that D_{jt}^{-6} is now an indicator for exactly 6 years prior to relocation.

Figure 2 presents the θ coefficients from this alternative specification. The θ coefficients are now consistently near zero prior to relocation, yet the postrelocation θ coefficients are near identical to the corresponding estimates from the baseline model. Hence, I conclude that differential pretrends cannot account for the sharp drop in establishment black share following relocation, either qualitatively or quantitatively.

In the online appendix, I also show that the change in the racial composition of employees at relocating establishments is not driven by coincident changes in the occupational composition of employees at those establishments. Black workers are substantially less likely to work the same job following its relocation to the suburbs.

While an establishment's drop in normalized black share following relocation is large, the average drop one would have predicted for these establishments using the cross-sectional relationship between establishment location and normalized black share is about 50% larger at 0.35.[18] This discrepancy may reflect some combination of the following: (1) the causal effect of location is heterogeneous, and relocating establishments are atypical[19]; (2) the β coefficients from equation (1) are not the causal effect of establishment location; and (3) the effect of an establishment *location* is not the same as the effect of an establishment *relocation*. The effects may not coincide if, for example, central city residents are more aware of job opportunities at establishments that have relocated to the suburbs than establishments that are already located in the suburbs. Nonetheless, the effect of relocation is substantial. Given that, as documented in section IV, the share of MSA jobs located in the central city declines by 10% per decade from 1970 to 2000, the estimated causal effect of establishment relocations suggests that job suburbanization will reduce the black share of the workforce by about 2.3% each decade.

IV. Job Suburbanization and Racial Inequality across Labor Markets

I have shown that conditional on job characteristics, black workers are less likely than white workers to work in suburbs, and this segregation persists despite widespread job suburbanization over time. This finding suggests that job suburbanization may reduce black labor market opportunities and increase racial labor market inequality. However, suburbanization may be offset by worker resorting across workplaces so that, at the market level, the effect of suburbanization on racial inequality is muted. In this section, I exploit variation across MSAs to estimate the local labor market-level relationship between job suburbanization and black employment rates and earnings using census data.

In analyzing labor market outcomes, I restrict attention to men and women between the ages of 24 and 63 who are non-Hispanic white or black. To measure employment, I use an indicator for whether an individual is currently labor market "active," meaning employed or in school.[20] In practice, this measure reflects employment because only a small fraction of individuals in my sample report being in school, and this does not differ significantly by race. For this reason, I use "active" and "employed" interchangeably.

Combining data from each census, I construct a synthetic panel by collapsing individuals into cells and merging cells across years. I group individuals by MSA of residence, gender, race, education, and cohort. These groups are indexed by g. I exclude those who are institutionalized because individuals in that population may not be residing in their relevant labor market. This may attenuate the relevant estimates below given that incarceration rates began to increase substantially in the mid-1970s and a large share of black adults was incarcerated, though the cohorts I focus on will have largely "aged-out" of criminal activity by 1980 (Western & Pettit, 2000). Patterns for women should be less susceptible to this issue given their relatively low incarceration rates. I divide the sample into three education groups: less than high school graduate, high school graduate, and college graduate. I also divide the sample into cohorts, those who are the following ages in 1970: 24–33, 34–43, and 44–53.[21] I group by cohort rather than age because the intention is to follow the same group of individuals from decade to decade. Of course, the composition of cells may change from decade to decade due to migration; I explore the role migration plays in the analysis in the online appendix. I restrict the analysis to cells with

[18] In particular, I estimate a variant of equation (1), allowing the β coefficient to vary by MSA, and then use the estimated model to predict the normalized black share for all establishments in the data based on the MSA and distance from the CBD alone. I then calculate the average change in predicted normalized black share following relocation, averaging across relocating establishments.

[19] Note that relocating establishments have relatively low black shares even *prior to* relocation.

[20] The results are similar if I use weeks worked as the employment measure.

[21] This corresponds to individuals born in 1917–1926, 1927–1936, and 1937–1946.

TABLE 2.—SAMPLE DESCRIPTIVE STATISTICS, CELL LEVEL

	Black				White			
	1970	1980	1990	2000	1970	1980	1990	2000
Share	0.132	0.138	0.126	0.143	0.868	0.862	0.874	0.857
Female	0.563	0.565	0.574	0.576	0.515	0.517	0.517	0.518
1917–1926	0.287	0.285	—	—	0.341	0.343	—	—
1927–1936	0.335	0.317	0.455	—	0.313	0.310	0.478	—
1937–1946	0.378	0.398	0.545	1.000	0.346	0.346	0.522	1.000
<HS Grad	0.559	0.487	0.410	0.329	0.289	0.241	0.173	0.125
HS Grad	0.402	0.452	0.501	0.569	0.541	0.556	0.565	0.568
College Grad	0.038	0.060	0.089	0.102	0.170	0.203	0.262	0.307
Active, Male	0.868	0.747	0.693	0.546	0.947	0.869	0.819	0.714
	(0.059)	(0.121)	(0.144)	(0.113)	(0.030)	(0.106)	(0.127)	(0.090)
Active, Female	0.568	0.594	0.616	0.489	0.479	0.554	0.623	0.565
	(0.117)	(0.155)	(0.173)	(0.123)	(0.080)	(0.122)	(0.156)	(0.100)
π^{cc}	0.529	0.489	0.460	0.423	0.525	0.489	0.459	0.422
	(0.111)	(0.125)	(0.156)	(0.169)	(0.112)	(0.120)	(0.146)	(0.160)
$\Delta\log(\pi^{cc})$	—	−0.094	−0.094	−0.110	—	−0.089	−0.089	−0.107
	—	(0.110)	(0.137)	(0.079)	—	(0.104)	(0.126)	(0.073)
Share Living in	0.795	0.734	0.678	0.617	0.393	0.324	0.292	0.270
Central City	(0.131)	(0.155)	(0.181)	(0.207)	(0.157)	(0.150)	(0.153)	(0.155)

This table includes 58 consistently identified MSAs with the largest black populations in 1970 and only cells with at least 25 individuals. Statistics are weighted by cell size. See section IV for further details on cell construction. "Active" refers to the share of a cell employed or in school. π^{cc} refers to the fraction of MSA jobs located in the central city. "Share Living in Central City" refers to the share of an entire racial group living in the central city of each cell's MSA, not the share of cell living in the central city. The latter cannot be identified in all years of the census data.

at least 25 observations and weight cells by their size. This leaves 1,607 cells in 1970 over 58 MSAs.

Table 2 presents summary statistics on the demographics and labor market outcomes for the synthetic cohorts. Black men are employed at lower rates than white men in 1970 and experience larger proportional declines in employment in the short run and long run. By contrast, black women are employed at higher rates than white women in 1970, though white women experience larger increases in employment rates over time. From 1970 to 2000, the share of the black population living in the central city declines by 22%, while the share of the white population living in the central city declines by 31%. The share of MSA jobs located in the central city declines by 9%–11% each decade.

A. *Empirical Strategy*

I test the spatial mismatch hypothesis by estimating the relationship between job suburbanization and black-white inequality in employment rates and earnings, exploiting variation across MSAs. In particular, I test whether black cohorts experience larger declines in employment rates and earnings relative to comparable white cohorts in MSAs that experience more job suburbanization. I estimate linear differences-in-differences models of the form

$$\Delta Y_{mg} = \alpha_g + \beta\Delta\log(\pi^{cc}_m) + f(Y^{t_0}_{mg})\gamma + \text{black}_g \times \left(\beta^B\Delta\log(\pi^{cc}_m) + f(Y^{t_0}_{mg})\gamma^B\right) + \epsilon_{mg}, \qquad (4)$$

where g indexes groups, m indexes markets, α_g are group fixed effects, and black_g is an indicator for a cell of black individuals. Y_g is either the log employment rate or log average annual earnings corresponding to a cell g.[22] In some specifications I do not condition on baseline employment rates (earnings), but in others I specify $f(\cdot)$ as a quadratic function. I include a control for a polynomial in baseline Y because employment or earnings growth may depend nonlinearly on baseline employment or earnings.[23] In some specifications I include MSA-by-education category fixed effects [and omit the collinear $\Delta\log(\pi^{cc}_m)$ term] to isolate within-skill group racial differences in outcomes. The coefficient β has the following elasticity interpretation: a 1% decline in the fraction of MSA jobs located in the central city is associated with a β% decline in cell employment. The coefficient β^B reflects the *relative* decline for black cells. I cluster standard errors at the MSA level.

Before describing the baseline results, I explore how job suburbanization relates to other baseline cell-level and MSA-level characteristics. One concern with the empirical strategy described here is that job suburbanization may occur in areas where employment is already declining, particularly for black workers. Though the 1960 census does not identify MSAs, for half the respondents in the 1970 census I can observe their employment status in 1965. To measure pretrends by cell, I use the change in employment rates by cell, assigning individuals to MSAs using their residence in 1970. I estimate models of the form

$$\Delta^{PRE}\log(\text{Emp. Rate})_{mg} = \alpha_g + \beta\Delta\log(\pi^{cc}_m) + \beta^B\text{black}_g \times \Delta\log(\pi^{cc}_m) + \epsilon_{mg}. \qquad (5)$$

[22] The results are similar if I use the employment rate rather than log employment rate as the outcome of interest.

[23] For example, because the cell employment rate is capped at 1, cells with high baseline employment rates have very little potential for growth relative to cells with lower baseline employment.

TABLE 3.—JOB SUBURBANIZATION AND BASELINE CELL AND MSA CHARACTERISTICS

Panel A: Employment (1965–1970) Outcome: Δ^{PRE} log (Emp. Rate)	1970–1980	1970–2000
$\Delta \log(\pi^{cc})$	0.083~	0.034
	(0.049)	(0.060)
$\Delta \log(\pi^{cc}) \times$ black	0.058	−0.040
	(0.040)	(0.027)
Group FEs	Yes	Yes
N Cells	1607	563
N MSAs	58	58

Panel B: 1970 Characteristics	1970–1980		1970–2000	
Outcome: $\Delta \log(\pi^{cc})$	δ	δ^B	δ	δ^B
Log Fraction Active (Group)	0.318~	−0.018	0.246	−0.088
	(0.160)	(0.095)	(0.300)	(0.209)
Log Earnings (Group)	−0.168	−0.056	−0.045	−0.281
	(0.098)	(0.071)	(0.229)	(0.196)
Fraction of Work in Central City (MSA),	−0.008	−0.002	0.072*	0.019
Standardized	(0.017)	(0.010)	(0.031)	(0.021)
Fraction Black (MSA),	−0.022	−0.020	−0.114**	−0.058~
Standardized	(0.019)	(0.014)	(0.039)	(0.033)
Dissimilarity Index (MSA),	−0.036**	−0.009	−0.098**	−0.016
Standardized	(0.012)	(0.011)	(0.026)	(0.018)
Violent Crime Rate (MSA),	−0.004	0.012	0.110	0.011
Standardized	(0.019)	(0.011)	(0.052)	(0.026)
Property Crime Rate (MSA),	−0.005	−0.003	−0.089*	0.006
Standardized	(0.017)	(0.012)	(0.036)	(0.025)
Group FEs	Yes		Yes	
N Cells	1564		547	
N MSAs	56		56	

Panel A displays estimates of equation (5). In panel A, the outcome is cell-level changes in employment rates from 1965 to 1970. Panel B displays estimates of equation (6). In this panel, the outcome is MSA job centralization. Columns 1 and 2 refer to centralization from 1970 to 1980; columns 3 and 4 refer to centralization from 1970 to 2000. The odd columns display the estimated δ coefficients from equation (6), the "main effects"; the even columns display the estimated δ^B coefficients, the black cell interactions. All models include group fixed effects (fixed effects for all combinations of cohort, education, sex, and race). Dissimilarity segregation indices are taken from Cutler et al. (1999). Standard errors are in parentheses, clustered at the MSA level. Regression models are weighted by cell size. ~Denotes statistical significance at the $p < 0.10$ level. *Denotes statistical significance at the $p < 0.05$ level. **Denotes statistical significance at the $p < 0.01$ level.

I estimate separate models for job suburbanization occurring over the short run (1970–1980) and long run (1970–2000). Table 3, panel A provides the results. I find that cells that experienced more job suburbanization from 1970 to 1980 had somewhat lower employment growth from 1965 to 1970. For suburbanization occurring from 1970 to 2000, the magnitude is even smaller and statistically insignificant. Importantly, the differences in trends between black and white cells are statistically insignificant in both cases and have opposite signs. Hence, black employment does not appear to be on a differential trend in suburbanizing MSAs.

A second concern is that job suburbanization is associated with other covariates that may influence labor market outcomes differently for black and white workers. Table 3, panel B provides coefficient estimates for models in the form

$$\Delta \log(\pi_m^{cc}) = \alpha_g + W_{mg}\delta + \text{black}_g \times W_{mg}\delta^B + \epsilon_{mg}, \qquad (6)$$

where W_m is a vector of cell-level and MSA-level covariates. I relate these correlates to short run and long run subsequent job suburbanization. I relate job suburbanization to the following baseline cell-level characteristics: log employment rate, log average earnings. I also include the following baseline MSA-level characteristics: fraction of jobs in the central city, black share of population, racial residential segregation, violent crime rate, and property crime rate. To measure residential segregation, I use the dissimilarity index constructed in Cutler et al. (1999). Data on reported crimes come from the FBI's Uniform Crime Reports. The UCR reports crimes per 100,000 for seven types of offenses: murder, rape, robbery, aggravated assault, burglary, larceny, and motor vehicle theft. I divide these seven offenses into two categories, violent and property crimes, and sum within these categories. I standardize the MSA-level covariates to have mean zero and standard deviation of one across MSAs.

From 1970 to 1980, I find residential segregation to be a significant predictor of job suburbanization across cells. Baseline employment rates are a marginal predictor. Notably, these relationships do not differ significantly for black cells. From 1970 to 2000, baseline residential segregation and the black share of the population are significant predictors of job suburbanization. This is consistent with research suggesting that black in-migration to central cities was a major cause of "white flight" (Boustan, 2010). The baseline fraction of jobs located in the central city is a marginal predictor.

B. *Baseline Estimates*

I estimate variants of equation (4) for three periods: 1970–1980, 1970–1990, and 1970–2000. Table 4 provides these baseline estimates. In the top panel, the outcome is log employment rate. In the bottom panel, the outcome is log average

TABLE 4.—JOB SUBURBANIZATION AND LABOR MARKET OUTCOMES

	1970–1980			1970–1990			1970–2000		
Outcome: Δ log (Emp. Rate)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\Delta \log(\pi^{cc})$	−0.030	−0.020		−0.023	−0.006		−0.039	−0.029	
	(0.051)	(0.057)		(0.071)	(0.049)		(0.048)	(0.037)	
$\Delta \log(\pi^{cc}) \times$ black	0.247^{**}	0.286^{**}	0.229^{**}	0.233^{**}	0.181^{**}	0.141^{**}	0.232^{**}	0.163^{**}	0.121^{**}
	(0.065)	(0.075)	(0.057)	(0.054)	(0.057)	(0.051)	(0.041)	(0.038)	(0.044)
Outcome: Δ log (Avg. Earnings)									
$\Delta \log(\pi^{cc})$	−0.030	−0.050		0.080	0.093		0.027	0.036	
	(0.066)	(0.075)		(0.101)	(0.105)		(0.062)	(0.066)	
$\Delta \log(\pi^{cc}) \times$ black	0.365^{**}	0.160	$0.140^{\sim}$	0.230^{*}	0.143	0.016	0.233^{**}	0.124^{*}	0.030
	(0.097)	(0.094)	(0.080)	(0.076)	(0.079)	(0.073)	(0.072)	(0.057)	(0.057)
Group FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
MSA × Education FEs	No	No	Yes	No	No	Yes	No	No	Yes
Baseline Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
N Cells	1,607	1,607	1,607	1,099	1,099	1,099	563	563	563
N MSAs	58	58	58	58	58	58	58	58	58

The table displays estimates of equation (4). Columns 1–3 refer to models covering 1970–1980, columns 4–6 refer to models covering 1970–1990, and columns 7–9 refer to models covering 1970–2000. All models include group fixed effects (fixed effects for all combinations of cohort, education, sex, and race). Columns 2, 3, 5, 6, 8, and 9 include a quadratic in log baseline employment rates (panel A) or log average earnings (panel B) interacted with race. Columns 3, 6, and 9 include MSA-by-education category fixed effects. In panel A, the outcome is changes in log employment rates; in panel B, the outcome is changes in log average earnings. All models are estimated at the cell level. Standard errors are in parentheses, clustered at the MSA level. Regression models are weighted by cell size. $^{\sim}$Denotes statistical significance at the $p < 0.10$ level. *Denotes statistical significance at the $p < 0.05$ level. **Denotes statistical significance at the $p < 0.01$ level.

earnings. In columns 1–3 the period is 1970–1980, in columns 4–6 the period is 1970–1990, and in columns 7–9 the period is 1970–2000. All columns include group fixed effects. All columns except 1, 4, and 7 include controls for a quadratic in baseline employment rates or log average earnings interacted with race. Columns 3, 6, and 9 include MSA-by-education fixed effects.

Across specifications and periods, changes in the spatial distribution of work have little association with white employment rates. The coefficient is generally small in magnitude and is statistically indistinguishable from zero at the 10% level. By contrast, job suburbanization is associated with declines in black employment, and the relationship is statistically significant at the 1% level across specifications. I first focus on specifications that do not include MSA-by-education fixed effects (all but columns 3, 6, and 9). Over the 1970s, the β^B coefficient of 0.25–0.29 implies that a 10% decline in the fraction of MSA jobs located in the central city—about the mean level experienced over this period—is associated with a 2.5%–2.9% decline in the black employment rate, relative to the white employment rate. The estimates for annual earnings have the same sign, but the coefficients are more dependent on the inclusion of baseline earnings and MSA-by-education fixed effects as controls.

From 1970 to 1990, a 10% decline in the fraction of jobs located in the central city is associated with a 1.8%–2.3% decrease in black relative employment. Over the full period, a 10% decline in the fraction of jobs located in the central city is associated with a 1.6%–2.3% decrease in black relative employment. The latter relationship is estimated using a single cohort of workers: those who were ages 24–33 in 1970 and 54–63 in 2000. Figure 3 displays this relationship graphically, plotting $\Delta_m \log(\pi^{cc})$ against changes in log black and white employment rates for *all* individuals in these cohorts pooled by MSAs. For black workers, there is a clear, positive relationship; for white workers, there is no clear relationship.

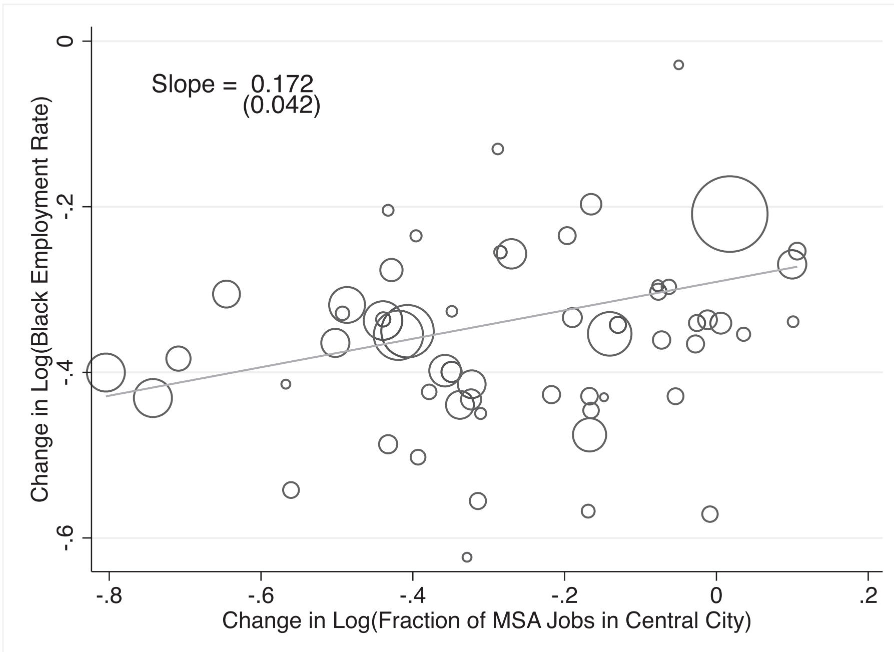
A potential explanation for the negative relationship between job suburbanization and black labor market outcomes is that it reflects racial differences in skill. Columns 3, 6, and 9 include MSA-by-education fixed effects so that β^B is identified by within-skill group variation. β^B decreases slightly in magnitude with the inclusion of these fixed effects, but the differences are not statistically significant. The relationship between job suburbanization and racial labor market inequality is a within-skill group phenomenon.

In online appendix table A4, I examine heterogeneity in the baseline estimates by education and gender. I focus on the 1970–2000 long difference. All models include a quadratic in baseline employment or earnings. For employment, the β^B coefficient is largest for high school dropouts but is also present for high school and college graduates. The coefficient is somewhat larger for women than men, though the difference is not statistically significant. This is striking given that men and women tend to work in very different industries and occupations over this period. For earnings, the relationship is present for women but not men.

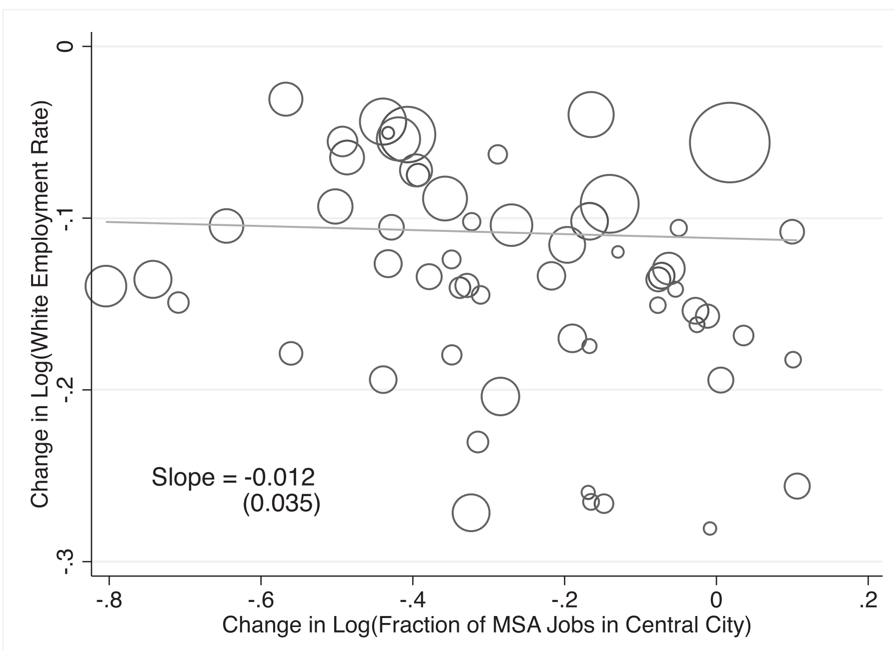
Given that the average value for $\Delta \log(\pi^{cc})$ is about −0.1 for each decade, the observed job suburbanization predicts a 1.6%–2.3% relative decrease in black employment rates. This decrease is comparable in magnitude to the decline in black share of the workforce implied by the establishment relocation event study estimates as described in section III, and it suggests that the resorting of black workers across workplaces did not significantly offset the effects of job suburbanization. To put this magnitude in perspective, note that from 1970 to 2000, the proportion of black men ages 24–63 living in the MSAs analyzed here that were employed or in school decreased from 85 percentage points (p.p.) to

FIGURE 3.—JOB SUBURBANIZATION AND CHANGES IN EMPLOYMENT RATES, 1970–2000

(A) Synthetic Panel, Black



(B) Synthetic Panel, White



This figure plots changes in black and white employment rates against job centralization across 58 MSAs from 1970 to 2000 for those born between 1937 and 1946 (ages 24–33 in 1970 and 54–63 in 2000). See section IVB for details on the construction of synthetic cohorts.

72 p.p. For white men, the proportion decreased from 92 p.p. to 87 p.p. Among women, the proportion employed or in school increased from 56 p.p. to 68 p.p. for black women and from 47 p.p. to 73 p.p. for white women. Hence, the employment rate for black men declined by about 10% more than the employment rate for white men, and the employment rate for white women increased by about 34% more than the employment rate for black women. β^B estimates imply that job suburbanization can explain about 50%–70% of the relative decline in black men's employment rates and 15–20% of the relative increase in white women's employment rates.

The evidence presented in table 4 and figure 3 suggests that job suburbanization caused substantial declines in black employment rates relative to white employment rates. However, there are three reasons to be cautious in assigning a causal interpretation. First, job suburbanization may be an endogenous response to unobserved, racially distinct labor supply shocks. In particular, job suburbanization may be concentrated in markets where black workers become relatively less productive so that black employment rates would have declined in those markets even if jobs had not moved to the suburbs. Second, suburbanization may be related to other changes in the local industry or occupation mix that would have otherwise portended reductions in black relative employment rates. Third, job suburbanization may be associated with migration patterns in a way that contaminates the synthetic panel design. I address each of these concerns below.

C. *Highways as a Source of Variation*

Unobserved shocks to the productivity of black workers could induce firms to migrate *and* produce black employment declines, even in the absence of a spatial mismatch mechanism. For example, the emergence of crack cocaine markets in the 1980s and 1990s could potentially explain both the deterioration of black labor supply and the relocation of employers from the central city (Evans et al., 2016).

To test whether the observed relationship between job suburbanization and declines in black relative employment is driven entirely by such productivity shocks, I require an instrument for job suburbanization that is plausibly orthogonal to such shocks. More generally, the ideal instrument would satisfy an exclusion restriction: it would only (potentially) affect black employment by changing the spatial distribution of work.

To instrument for job suburbanization, I exploit a previously used source of variation in suburbanization: the interstate highway system (Baum-Snow, 2007; Michaels, 2008). Highways can potentially increase suburbanization through several mechanisms. First, they decrease transportation costs for both firms and households. For firms, highways make physical proximity to other transportation hubs (e.g., ports and rail stations) and upstream or downstream firms less important, allowing them to take advantage of cheaper land and other suburban amenities. Michaels (2008) argues that as highway construction was nearing completion, trucks became the primary mode of shipping goods within the United States. For households, highways reduce the costs of commuting to central work and access to other central city amenities from a suburban residence. These direct effects on transportation costs may also have feedback effects. By increasing the number of firms and households in the suburbs, they make these areas more attractive for other firms and households. Firms may then follow workers moving to the suburbs and achieve agglomeration economies there.

Baum-Snow (2007) documents that the U.S. interstate highway system played an important role in postwar residential suburbanization. With nearly all construction occurring between 1956 and 1980, the interstate highway system would ultimately span over 40,000 miles. The highway system was originally designed to connect major metropolitan areas, serve U.S. national defense, and connect major routes in Canada and Mexico. Using plausibly exogenous variation in highway construction across MSAs, Baum-Snow (2007) finds that one new highway passing through a central city reduces its population from 1950 to 1990 by about 18%. These effects are substantial: they imply that the interstate highway system accounts for about one-third of the decline in central city population relative to total MSA growth over this period. Using a similar identification strategy, Baum-Snow (2020) finds that highways also caused substantial job suburbanization.

Highways are an imperfect instrument for job suburbanization: research has documented that highways have a variety of effects on the labor market, and it is possible that they affect black relative employment through channels other than the spatial distribution of work.[24] However, in the online appendix, I provide evidence that highways do not appear to affect labor demand for worker skills in a way that would disproportionately affect black workers.

Highways also affect residential neighborhoods. In the online appendix, I show that, in the sample of MSAs examined here, highway rays predict the suburbanization of white households at a magnitude consistent with Baum-Snow (2007) but not the location of black households. This disparity is consistent with a central premise of the spatial mismatch hypothesis: black households faced significant additional barriers to suburban residence over this period. Accordingly, highway rays also increase subsequent segregation, with each ray predicting a 0.01, 0.045, and 0.053 unit increase in a city's dissimilarity index by 1980, 1990, and 2000.

Building the interstate highway system also forced the destruction of neighborhoods and displacement of households, particularly in central cities. There is evidence that black households faced a disproportionate share of displacements and that local planners exploited interstate construction as an opportunity to eliminate poor, "blighted," and often black communities (Rose & Mohl, 2012). This suggests another channel through which the interstate highways system may affect black relative employment rates, potentially violating the exclusion restriction. However, more than 90% of interstate-central city intersections in the MSAs I study were already built by 1970. Hence, the effects of neighborhood clearances would likely already be reflected in baseline labor market outcomes. For the analysis below, my measure of MSA highway exposure is the stock of interstate highway rays emanating from the central city in 1970.

Are highways endogenous to labor supply shocks? A potential concern with exploiting variation derived from the

[24]Michaels (2008) finds that highways increased trade for exposed rural counties and raised the relative demand for skilled manufacturing workers in skill-abundant rural counties while reducing it elsewhere. Duranton et al. (2014) find that highways lead cities to increase the weight of their exports and specialize in sectors producing heavy goods. Duraton and Turner (2012) find that interstate highways increased MSA growth from 1983 to 2003.

interstate highway system is that highway assignment may be determined endogenously. As Baum-Snow (2007) notes, the interstate highway system was likely designed in part to facilitate local commuting and local economic development in particular regions. Though less plausible, the highway system might have also been designed accounting for productivity shocks particular to black workers in the 1970s. To deal with this, I instrument for realized highway construction using the 1947 federal interstate highway plan as in Baum-Snow (2007).

In 1937, the Franklin D. Roosevelt administration began to plan an interstate highway system. In their recommended highway plan, the National Interregional Highway Committee considered the distribution of population, manufacturing activity, agricultural production, the location of post-World War II employment, a strategic highway network drawn up by the War Department in 1941, military and naval establishment locations, and interregional traffic demand, in that order. This led to the Federal Highway Act of 1944, which instructed the roads commissioner to develop an initial plan for a national interstate highway system. As specified by the legislation, the highways in the planned system were to be "so located as to connect by routes as direct as practicable, the principal metropolitan areas, cities, and industrial centers, to serve the national defense, and to connect at suitable border points with routes of continental importance in the Dominion of Canada and the Republic of Mexico. . ." (as cited in Baum-Snow, 2007). Importantly, the legislation makes no mention of local commuting or local economic development. The final plan produced under this act, approved in 1947, is presented in online appendix figure A2.

Major funding for the interstate highway system began with the Federal Aid Highway Act of 1956. The 1956 Federal Aid Highway Act expanded the 1947 plan and committed the federal government to pay 90% of the cost of construction. In particular, the 1956 plan incorporated additional highways that were designed for local purposes like commuting and development. Therefore, in some specifications I instrument for actual highway rays using highway rays planned in 1947. The first-stage t-statistic is in excess of 5.

Empirical strategy and results. I explore the relationship between highways, job suburbanization, and employment rates by race. In all subsequent analyses, my measure of highway exposure is the number of interstate highway rays emanating from the central city in 1970. I estimate specifications of the form

$$\begin{aligned}\Delta Y_{mg} = {} & \alpha_g + \gamma_1 rays_m^{1970} + \gamma_2 radius_m + f(emp_{mg}^{1970}) \\ & + X_m\gamma_3 + \text{black}_g \times \left(\gamma_1^B rays_m^{1970} + \gamma_2^B radius_m \right. \\ & \left. + \; + X_m\gamma_3^B\right) + \epsilon_{mg}, \end{aligned} \tag{7}$$

where $rays_m^{1970}$ denotes the number highway rays emanating from the central city of MSA m in 1970, and $radius_m$ is the radius of the central city, a key control in the analysis of Baum-Snow (2007). Intuitively, one must control for the central city radius because it determines the extent to which sprawl is reflected in suburbanization measures, and highways are more likely to travel through central cities that are physically larger. Note that the average number of central city highway rays in 1970, weighted by the sample population size, is 3.9; the unweighted average is 3.[25]

As in table 3, I begin by relating highways to preperiod changes in employment rates and baseline cell and MSA characteristics. The results are presented in online appendix table A5. There is no relationship between an MSA's highway stock in 1970 and cell-level changes in employment rates from 1965 to 1970. Notably, highway assignment has little relationship with an MSA's racial composition in 1970. Similarly, Baum-Snow (2007) finds that planned and actual highway construction has little relationship with an MSA's racial composition in 1950.

Table 5, panel A presents estimates of equation (7) where the outcome is $\Delta_m \log(\pi^{cc})$. I find that the stock of highways in 1970 predicts job suburbanization thereafter. In odd specifications I use actual interstate highway rays constructed as the explanatory variable of interest; in even specifications I instrument for highways constructed using highway rays included in the 1947 plan. In all columns, t_0 is 1970, while t_1 is 1980 in columns 1 and 2, 1990 in columns 4 and 5, and 2000 in columns 7 and 8. In the baseline OLS specification, one highway ray emanating from the central city is associated with a 3.5%, 5.8%, and 7.2% decrease in the fraction of MSA jobs located in the central city by 1980, 1990, and 2000. When I instrument for highways, the point estimates increase somewhat, particularly when weighting by black population, though they are less precise. The estimated relationship between interstate highways and job suburbanization is comparable in magnitude to the relationship documented in Baum-Snow (2020), who finds that each new radial highway decentralized 4%–6% of jobs from 1960 to 2000 in a larger set of MSAs.

Next, I estimate the reduced-form effect of highways on black relative employment and earnings. I estimate specifications of the same form as equation (7), where Y is now log employment rate or log earnings. Panels B and C of table 5 present the estimates. Columns 1, 4, and 7 are OLS models, while in the remaining specifications, I instrument for actual highway rays in 1970 using planned highway rays. All specifications include a quadratic in baseline employment. Columns 3, 6, and 9 include MSA-by-education fixed effects. The pattern of coefficients is consistent across specifications and outcomes. While highways predict increases in the employment rates and earnings of whites, consistent with Duraton and Turner (2012), they predict relative decreases in these labor market outcomes for black adults. For the employment outcomes, the magnitudes of the coefficients are stable across specifications.

[25] Note that if a highway passes through a central city, this counts as two rays.

TABLE 5.—HIGHWAYS, JOB SUBURBANIZATION, AND LABOR MARKET OUTCOMES

	1970–1980			1970–1990			1970–2000		
	OLS	IV	IV	OLS	IV	IV	OLS	IV	IV
Outcome: Δ log(π^{cc})	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Rays^{1970}	−0.035**	−0.037*		−0.058**	−0.079**		−0.072**	−0.110**	
	(0.010)	(0.017)		(0.015)	(0.024)		(0.020)	(0.030)	
Rays^{1970} × black	−0.005	−0.016		−0.005	−0.024		−0.007	−0.027	
	(0.006)	(0.009)		(0.007)	(0.013)		(0.009)	(0.018)	
Outcome: Δ log (Emp. Rate)									
Rays^{1970}	0.008*	0.0110*		0.009*	0.015*		0.012**	0.020*	
	(0.004)	(0.005)		(0.004)	(0.006)		(0.004)	(0.008)	
Rays^{1970} × black	−0.022**	−0.021*	−0.015*	−0.018*	−0.018	−0.012	−0.019**	−0.028**	−0.021**
	(0.005)	(0.009)	(0.007)	(0.007)	(0.011)	(0.009)	(0.007)	(0.010)	(0.008)
β, IV	−0.235~	−0.295		−0.156~	−0.188~		−0.156~	−0.180~	
	(0.132)	(0.229)		(0.079)	(0.106)		(0.80)	(0.096)	
β^B, IV	0.575**	0.486*		0.288*	0.220*		0.249**	0.242**	
	(0.198)	(0.224)		(0.121)	(0.105)		(0.094)	(0.087)	
Outcome: Δ log (Earnings)									
Rays^{1970}	0.009~	0.021**		0.011	0.006		0.021*	0.015	
	(0.005)	(0.007)		(0.007)	(0.012)		(0.010)	(0.014)	
Rays^{1970} × black	−0.020**	−0.022*	−0.021*	−0.020	−0.015	0.003	−0.016	−0.022	−0.008
	(0.006)	(0.011)	(0.009)	(0.011)	(0.015)	(0.013)	(0.011)	(0.018)	(0.012)
β, IV	−0.285	−0.616		−0.192	−0.074		−0.222~	−0.143	
	(0.195)	(0.394)		(0.147)	(0.166)		(0.127)	(0.150)	
β^B, IV	0.586**	0.630		0.364	0.172		0.239	0.192	
	(0.221)	(0.393)		(0.233)	(0.172)		(0.173)	(0.150)	
Group FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
MSA × Education FEs	No	No	Yes	No	No	Yes	No	No	Yes
Baseline Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
CC Radius Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N Cells	1559	1559	1559	1066	1066	1066	545	545	545
N MSAs	56	56	56	56	56	56	56	56	56

The table displays estimates of equation (7). Columns 1–3 refer to models covering 1970–1980, columns 4–6 refer to models covering 1970–1990, and columns 7–9 refer to models covering 1970–2000. In columns 2, 3, 5, 6, 8, and 9, actual highway rays in 1970 are instrumented with planned highways. All models include group fixed effects (fixed effects for all combinations of cohort, education, sex, and race), a quadratic in log baseline employment (panels A and B) or log average earnings (panel C) interacted with race, and central city radius interacted with race. Columns 3, 6, and 9 include MSA-by-education fixed effects. In panel A, the outcome is changes in log fraction of MSA jobs located in the central city; in panel B, the outcome is changes in log employment rates; in panel C, the outcome is changes in log earnings. "β, IV" and "β^B, IV" are the implied IV estimates for β and β^B, where highway rays or planned rays are instrumented for $\Delta \log(\pi^{cc})$. All models are estimated at the cell level. Two MSAs are not included in the highway data (Baum-Snow, 2007): Jackson, MS and West Palm Beach, FL. Standard errors are in parentheses, clustered at the MSA level. Regression models are weighted by cell size. ~Denotes statistical significance at the $p < 0.10$ level. *Denotes statistical significance at the $p < 0.05$ level. **Denotes statistical significance at the $p < 0.01$ level.

In the OLS models, each additional highway ray predicts a 1% increase in white employment rates but about a 2% relative decline in black employment. Coefficients from the IV models are similar. Strikingly, while a racial gap in employment rates emerges from 1970 to 1980, it remains roughly constant through 2000. The γ_1^B coefficient estimate does not change significantly with the inclusion of MSA-by-education fixed effects, indicating the relationship between highways and black relative employment rates holds within skill groups. As in section IVB, the magnitude of the coefficients in the earnings models is sensitive to whether I control for baseline earnings.

Highway construction causes job suburbanization and increases the gap in employment rates between white and black workers. In table 5, I also include IV estimates for β and β^B, the slope coefficients for $\Delta \log(\pi^{cc})$ in equation (4), where I instrument for $\Delta \log(\pi^{cc})$ using highway rays or planned rays.[26] In specifications corresponding to columns 1, 3, 5, and 6, the first-stage (panel A of table 5) Angrist-Pischke F-statistics exceed 10, allaying concerns over weak identification in these models (Angrist and Pischke, 2009).[27] columns 5 and 6 indicate that for every 10% decrease in the fraction of jobs located in the central city induced by highways from 1970 to 2000, black employment rates decline by 2.4%–2.5% relative to white employment rates. While job suburbanization might not be the only mechanism through which highways affect relative employment rates, labor supply shocks are unlikely to be an alternative, and changes in labor demand are also unlikely to be a significant channel. Yet the relationship between job suburbanization and black employment implied here is comparable to the baseline estimates from section IVB,[28] corroborating a causal

[26]More specifically, I estimate variants of equation (7) where I replace $rays_m^{1970}$ with $\Delta \log(\pi^{cc})$ and instrument for $\Delta \log(\pi_m^{cc})$ and $\text{black}_g \times \Delta \log(\pi_m^{cc})$ using $rays_m^{1970}$ and $\text{black}_g \times rays_m^{1970}$. β and β^B correspond to the coefficients on $\Delta \log(\pi_m^{cc})$ and $\text{black}_g \times \Delta \log(\pi_m^{cc})$.

[27]In the specification corresponding to column 4, the Angrist-Pischke F-statistics associated with Rays^{1970} and Rays^{1970} × black are 8.63 and 13.2. In the specification corresponding to column 2, both are below 10.

[28]The IV estimates are generally larger in magnitude but also less precise, and so they are statistically indistinguishable from baseline estimates.

interpretation of the relationship between job suburbanization and black employment.

D. *Additional Robustness Checks*

In the online appendix, I consider and rule out two alternative interpretations for the relationship between job suburbanization and black-white differences in employment rates documented above. First, I explore whether job suburbanization, including interstate highway-induced suburbanization, is associated with changes in the types of work performed in the labor market. Specifically, I test whether MSAs that experience greater job suburbanization also experience relative growth in industries or occupations that disproportionately employed white workers in 1970. I conclude that the causal effect of interstate highways on racial differences in employment rates is not driven by changes in sector composition.

Second, I test whether the relationship between job suburbanization and black-white differences in employment rates is driven by the endogenous migration of households. Due to data limitations, I conduct the main analysis using a synthetic panel rather than a true panel of individuals. For the same reason that residential sorting is a concern for any cross-sectional analysis, the endogenous migration of households to and from MSAs may introduce a compositional bias in the synthetic panel analysis. The productivity of migrants to and from an MSA may vary systematically with job suburbanization so that the correlation between job suburbanization and employment rate changes may reflect in part the changing composition of cells rather than within-person changes. Using a true panel of individuals from 1975 to 1980, I find that endogenous migration is not a first-order concern, and the relationship between job suburbanization and racial differences in employment rates is similar whether I use a true panel or a synthetic panel.

V. Mechanisms of Spatial Mismatch

This paper shows that, consistent with the spatial mismatch hypothesis, relocating a job to the suburbs reduces the likelihood that the worker holding that job is black, and that job suburbanization reduced black employment rates from 1970 to 2000. These findings raise a natural question: why? For example, why don't black workers follow jobs to the suburbs? And why don't more firms locate in the central city to employ displaced black workers? While a complete exploration of the mechanisms that generate a link between job suburbanization and racial inequality in the labor market is beyond the scope of this paper, this section summarizes potential mechanisms and the necessary ingredients for the spatial mismatch hypothesis to hold.

The spatial mismatch hypothesis predicts that the suburbanization of work reduces black relative employment via spatial frictions in the labor and housing markets. Coulson et al. (2001) develop a general equilibrium spatial search model that identifies the conditions necessary to generate higher unemployment rates in the central city due to spatial frictions. The model can be easily adapted to identify the conditions under which job suburbanization reduces black employment.

In the Coulson et al. (2001) framework, the following conditions are sufficient for spatial mismatch to emerge in equilibrium (Johnson, 2006):

1. Commuting or job search costs (as a function of distance) are nontrivial.
2. Black households are relatively constrained to residing in the central city.
3. Firms face higher (nonlabor) costs (either fixed or production costs) in the central city than in the suburbs.

Conditions 1 and 2 are essential because black workers must find it relatively difficult to work in the suburbs. With no search or commuting costs, the spatial distribution of work would have no bearing on labor market outcomes. The steep and accumulating declines in black employment following establishment relocations to the suburbs indicates that the combination of commuting and search costs is substantial. Even with search and commuting costs, if black and white households found it equally costly to reside in the suburbs, black households would be equally able to follow work to the suburbs. Indeed, there is evidence that black households faced relatively high barriers to suburban residence over this period due to discrimination in housing and mortgage markets (Yinger, 1986, 1995).

In addition, given their low levels of wealth (Blau & Graham, 1990; Barsky et al., 2002), black households may be more likely to face binding liquidity constraints in securing suburban housing, where land use zoning regulations may limit the supply of relatively low-cost housing (Rothwell & Massey, 2009; Rothwell, 2011). Finally, the continued concentration of black households in the central city may be market-driven if black and white households differ in their preferences over neighborhood attributes, including racial composition. For example, black households may prefer to live in racially diverse or predominantly black neighborhoods, which until recently were rare in the suburbs (Bayer et al., 2014).

Even with constraints on residential mobility and nontrivial search and commuting costs, the mobility of firms will tend to equalize employment opportunities across the central city and suburbs. If productive workers remain in the central city, competitive forces should induce the entry of a sufficient number of firms to absorb that labor until subsequent entry remains unprofitable. For the predictions of the spatial mismatch hypothesis to hold, it is critical that job suburbanization be driven in part by declines in the relative costs of operating in the suburbs faced by firms. This makes suburban entry more attractive, even ignoring concerns about labor accessibility. Holding central city labor productivity constant, central city entry becomes relatively less profitable, increasing the relative unemployment rate in the central city. Indeed, there is evidence that a number of nonlabor factors induced firm entry in the suburbs by reducing the relative fixed and production costs associated with operating there. These include innovations in transportation and transportation infrastructure, lower suburban land costs, and the formation of agglomeration economies in the suburbs (Glaeser & Kahn, 2001).

In addition, labor suburbanization can also reduce the effective productivity of workers who remain in the central city if complementarities exist between types of labor in the production function, production exhibits increasing returns to scale, or product demand is local. For example, if the presence of skilled managers in a plant increases the productivity of unskilled workers, then the migration of skilled labor out of the central city may reduce the effective productivity of the unskilled workers who remain.

VI. Discussion

For several decades, spatial mismatch has been a commonly cited cause of persistently high black unemployment. In this paper, I describe the process of job suburbanization and estimate its effects on racial labor market inequality from 1970 to 2000. I find that, conditional on job characteristics, black workers are less likely than white workers to work in suburbs, and this segregation persists despite widespread job suburbanization. Exploiting variation in suburbanization across local labor markets, I find that for every 10% decrease in the fraction of MSA jobs located in the central city, black relative employment rates declined by 1.6%–2.5%, with IV estimates derived from nationally planned highways at the high end of this range. This relationship holds within observable skill groups and for both men and women. Relative earnings declined by up to 2.4%, though these estimates are less stable. Conversely, job suburbanization is not related to other significant structural changes in the labor market that would portend changes in black relative employment rates. Estimates imply that job suburbanization can explain more than half of the relative decline in black men's employment rates (relative to white men's) and 15%–20% of the relative increase in white women's employment rates (relative to black women's) from 1970 to 2000.

Since 2000, the share of younger college graduates living in large central cities and high-wage jobs located in the central city has increased significantly (Couture & Handbury, 2020). The gap in employment rates between black men and women living in the central city and suburbs has remained stable over this period.[29] What effect this urban resurgence has on the spatial distribution of jobs more broadly and on racial inequality is an interesting question for future research.

[29] See the online appendix for additional details.

REFERENCES

Andersson, Fredrik, John C. Haltiwanger, Mark J. Kutzbach, Henry O. Pollakowski, and Daniel H. Weinberg, "Job Displacement and the Duration of Joblessness: The Role of Spatial Mismatch," this REVIEW C:2 (May 2018), 203–218.

Angrist, Joshua D., and Jörn-Steffen Pischke, *Mostly Harmless Econometrics: An Empiricist's Companion* (Princeton, NJ: Princeton University Press, 2009).

Barsky, Robert, John Bound, Kerwin Charles, and Joseph Lupton, "Accounting for the Black-White Wealth Gap: A Nonparametric Approach," *Journal of the American Statistical Association* 97:459 (September 2002), 663–673. 10.1198/016214502388618401

Baum-Snow, Nathaniel, "Did Highways Cause Suburbanization?" *Quarterly Journal of Economics* 122:2 (2007), 775–805. 10.1162/qjec.122.2.775

——— "Urban Transport Expansions and Changes in the Spatial Structure of US Cities: Implications for Productivity and Welfare," this REVIEW 102:5 (2020), 929–945.

Bayer, Patrick, Hanming Fang, and Robert McMillan, "Separate When Equal? Racial Inequality and Residential Segregation," *Journal of Urban Economics* 82 (July 2014), 32–48. 10.1016/j.jue.2014.05.002

Blau, Francine D., and John W. Graham, "Black-White Differences in Wealth and Asset Composition," *Quarterly Journal of Economics* 105:2 (1990), 321–339. 10.2307/2937789

Boustan, Leah, "Was Postwar Suburbanization 'White Flight'? Evidence from the Black Migration," *Quarterly Journal* 125:1 (February 2010), 417–443.

Carrington, William J., Kristin McCue, and Brooks Pierce, "Using Establishment Size to Measure the Impact of Title VII and Affirmative Action," *Journal of Human Resources* 35:3 (Summer 2000), 503–523. 10.2307/146390

Cattaneo, Matias, Richard Crump, Max Farrell, and Yingjie Feng, "On Binscatter," FRB of New York Staff Report No. 881 (2019).

Coulson, Edward, Derek Laing, and Paul Wang, "Spatial Mismatch in Search Equilibrium," *Journal of Labor Economics* 19:4 (October 2001), 949–972. 10.1086/322824

Couture, Victor, and Jessie Handbury, "Urban Revival in America," *Journal of Urban Economics* 119 (September 2020), 103267. 10.1016/j.jue.2020.103267

Cutler, David M., Edward L. Glaeser, and Jacob L. Vigdor, "The Rise and Decline of the American Ghetto," *Journal of Political Economy* 107:3 (June 1999), 455–506. 10.1086/250069

Duranton, Giles, Peter M. Morrow, and Matthew A. Turner, "Roads and Trade: Evidence from the US," *Review of Economic Studies* 81:2 (April 2014), 681–724. 10.1093/restud/rdt039

Duraton, Giles, and Matt Turner, "Urban Growth and Transformation," *Review of Economic Studies* 79:4 (October 2012), 1407–1440. 10.1093/restud/rds010

Ellwood, David T., "The Spatial Mismatch Hypothesis: Are There Teenage Jobs Missing in the Ghetto?" in Richard B. Freeman and Harry J. Holzer, eds., *The Black Youth Employment Crisis* (Chicago: University of Chicago Press, 1986).

Evans, William N., Craig Garthwaite, and Timothy J. Moore, "The White/Black Educational Gap, Stalled Progress, and the Long-Term Consequences of the Emergence of Crack Cocaine Markets," this REVIEW 98:5 (December 2016), 832–847.

Fairlie, Robert W., and William A. Sundstrom, "The Emergence, Persistence, and Recent Widening of the Racial Unemployment Gap," *Industrial and Labor Relations Review* 52:2 (January 1999), 252–270. 10.1177/001979399905200206

Fernandez, Roberto M., "Race, Spatial Mismatch, and Job Accessibility: Evidence from a Plant Relocation," *Social Science Research* 37:3 (September 2008), 953–975. 10.1016/j.ssresearch.2008.03.006

Glaeser, Ed, and Matthew Kahn, "Decentralized Employment and the Transformation of the American City," *Brookings-Wharton Papers on Urban Affairs* 2 (2001).

Holian, Matthew J., and Matthew E. Kahn, *The Impact of Center City Economic and Cultural Vibrancy on Greenhouse Gas Emissions from Transportation* (San Jose: MTI Publications, 2012).

Holzer, Harry J., and Keith R. Ihlanfeldt, "Spatial Factors and the Employment of Blacks at the Firm Level," *New England Economic Review* 65 (May/June 1996).

Ihlanfeldt, Keith R., and David L. Sjoquist, "The Spatial Mismatch Hypothesis: A Review of Recent Studies and Their Implications for Welfare Reform," *Housing Policy Debate* 9:4 (1998), 849–892. 10.1080/10511482.1998.9521321

Johnson, Rucker, "Landing a Job in Urban Space: the Extent and Effects of Spatial Mismatch," *Regional Science and Urban Economics* 36 (2006), 331–372. 10.1016/j.regsciurbeco.2005.11.002

Kain, John F., "Housing Segregation, Negro Employment and Metropolitan Decentralization," *Quarterly Journal of Economics* 82:3 (1968), 175–197. 10.2307/1885893

Michaels, Guy, "The Effect of Trade on the Demand for Skill: Evidence from the Interstate Highway System," this REVIEW 90:4 (2008), 683–701.

Mouw, Ted, "Job Relocation and the Racial Gap in Unemployment in Detroit and Chicago 1980–1990: A Fixed-Effects Estimate of the Spatial Mismatch Hypothesis," *American Sociological Review* 65 (2000), 730–753. 10.2307/2657544

Raphael, Steven, "The Spatial Mismatch Hypothesis and Black Youth Joblessness: Evidence from the San Francisco Bay Area," *Journal of Urban Economics* 43 (1998), 79–111. 10.1006/juec.1997.2039

Robinson, Corre L., Tiffany Taylor, Donald Tomaskovic-Devey, Catherine Zimmer, and Matthew W. Irvin, "Studying Race or Ethnic and Sex Segregation at the Establishment Level: Methodological Issues and Substantive Opportunities Using EEO-1 Reports," *Work and Occupations* 32:1 (February 2005), 5–38. 10.1177/0730888404272008

Rose, Mark H., and Raymond A. Mohl, *Interstate: Highway Politics and Policy since 1939* (Knoxville: The University of Tennessee Press, 2012).

Rothstein, Richard, *The Color of Law: A Forgotten History of How Our Government Segregated America*, 1st ed. (London: Liveright Publishing Corporation, 2017).

Rothwell, Jonathan T., "Racial Enclaves and Density Zoning: The Institutionalized Segregation of Racial Minorities in the United States," *American Law and Economics Review* 13:1 (Spring 2011), 290–358. 10.1093/aler/ahq015

Rothwell, Jonathan T., and Douglas S. Massey, "The Effect of Density Zoning on Racial Segregation in U.S. Urban Areas," *Urban Affairs Review* 44:6 (2009), 779–806. 10.1177/1078087409334163

Ruggles, Steven, J. Trent Alexander, Katie Genadek, Ronald Goeken, Matthew B. Schroeder, and Matthew Sobek, Integrated Public Use Microdata Series: Version 5.0 [Machine-readable database], Vol. 42 (Minneapolis: University of Minnesota Press, 2010).

Stoll, Michael A., "Job Sprawl, Spatial Mismatch, and Black Employment Disadvantage," *Journal of Policy Analysis and Management* 25:4 (2006), 827–854. 10.1002/pam.20210

Weinberg, Bruce A., Patricia B. Reagan, and Jeffrey J. Yankow, "Do Neighborhoods Affect Hours Worked? Evidence from Longitudinal Data," *Journal of Labor Economics* 22:4 (October 2004), 891–924. 10.1086/423158

Western, Bruce, and Becky Pettit, "Incarceration and Racial Inequality in Men's Employment," *Industrial and Labor Relations Review* 54:1 (October 2000), 3–16. 10.1177/001979390005400101

Wilson, William Julius, *The Truly Disadvantaged: The Inner City, the Underclass and Public Policy* (Chicago: University of Chicago Press, 1987).

——— *When Work Disappears: The World of the New Urban Poor* (New York: Alfred A. Knopf, 1996).

Yinger, John, "Measuring Racial Discrimination with Fair Housing Audits: Caught in the Act," *American Economic Review* 76:5 (December 1986), 881–893.

——— *Closed Doors, Opportunities Lost: the Continuing Costs and Housing Discrimination* (Manhattan, NY: Russell Sage Foundation, 1995).

Zax, Jeffery S., and John F. Kain, "Moving to the Suburbs: Do Relocating Companies Leave Their Black Employees Behind?" *Journal of Labor Economics* 14:3 (1996), 473–504. 10.1086/209819