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Race, space, and unemployment duration

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Abstract

This paper employs a special geo-coded tabulation of the US Panel Study of Income Dynamics (PSID) to examine the link between racial differences in residential location, or residential segregation, and racial disparities in unemployment durations. For a sample of 1098 workers whose employment was terminated during the years 1990 through 1992, we estimate continuous time duration models explaining unemployment durations for recently unemployed individuals as a function of personal and household characteristics affecting worker productivity, residential location characteristics at the time of job termination, and measures of job accessibility. We find that residential segregation affects racial differences in unemployment durations by exacerbating racial differences in job accessibility and neighborhood peer effects.

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1. Introduction

Recent nationwide statistics point to persistent racial gaps in employment outcomes. African American adults are more than twice as likely as white adults to be unemployed, according to the 2000 Census of Population and Housing (SF3). These racial gaps have remained unchanged over the last several decades [22]. Not only are African Americans more likely to be out of the labor force, but also among those workers who are unemployed, African Americans are more likely to remain unemployed for a longer period of time. According to the 2003 Current Population Survey, the median unemployment duration for African American workers over the age of 16 (12.9 weeks) is several weeks longer than the median unemployment duration for white workers (9.4 weeks).

As Holzer [5] and Clark and Summers [1] suggest, the link between the *incidence* of unemployment and the *duration* of unemployment for African Americans reflects persistent racial barriers to gaining rather than retaining employment, including labor market discrimination and housing market discrimination in areas experiencing job growth. This research provides a much needed perspective on the spatial factors contributing to the racial gap in unemployment durations. In particular, we ask the following questions: To what extent can black–white differences in the transition out of unemployment be explained by black–white differences in residential location, or residential segregation? What is the nature of the connection between segregation and unemployment durations? Finally, how important is residential location compared to automobile ownership status?

Our research is unique in its examination of the neighborhood-level influences on racial differences in the duration of unemployment for a national sample of US workers. This emphasis on employment durations is important, because while short-term unemployment is often caused by voluntary job-turnovers, long-term unemployment reflects fundamental structural impediments to obtaining work. Furthermore, the familial impacts of long-term unemployment are often more severe, especially in the wake of recent welfare reform initiatives which limit the duration of welfare assistance payments. Although a few studies have examined the link between residential location and racial differences in unemployment duration, ours is the first national study to simultaneously control for neighborhood peer effects, job accessibility, vehicle ownership status, along with a variety of other household and worker characteristics affecting unemployment durations. We also employ more precise measures of job accessibility than other multiple-MSA studies examining the link between residential location and racial differences in employment outcomes.

2. Background

In the standard spatial job search model (Rogers [16], Holzer et al. [6]), unemployed workers at time t choose between continued job search and accepting a wage offer. The probability (P_t) of accepting a wage offer and leaving unemployment at time t is equal to:

$$P_t = \alpha(e)\big[1 - F\big(w^*(c_{\mathrm{s}}, c_{\mathrm{c}})\big)\big], \tag{1}$$

where $\alpha(-)$ is the job offer arrival rate, e is the effort expended in job search, c_{s} is the cost of searching for a job, c_{c} is the cost of commuting to an accepted job, $F(w)$ is the cumulative distribution of job offers, and $[1 - F(w^*)]$ is the probability of accepting an offer that equals or exceeds the reservation wage, w^*.

One channel through which residential segregation may affect employment probabilities is via neighborhood influences on job search effort. We examine the impact of such neighborhood influences by examining how e affects P_t, holding reservation wages constant. Racial gaps in income and wealth accumulation suggest that blacks living in racially segregated neighborhoods are more likely to be surrounded by neighbors living below the poverty line. If black workers living in segregated neighborhoods have fewer opportunities to engage in local social interactions with employed middle to upper income adults, as suggested by Wilson [25], the absence of such peer influences may reduce incentives to search for employment. Furthermore, higher crime rates in high-poverty neighborhoods may reduce incentives to search for employment due to fear of being victimized on the way to work [7]. Assuming that search effort does not affect reservation wages, the effect of a change in search effort on the probability of leaving unemployment equals:

$$\frac{\mathrm{d}P_t}{\mathrm{d}e} = \frac{\mathrm{d}\alpha}{\mathrm{d}e}\bigl[1 - F\bigl(w^*(c_\mathrm{s}, c_\mathrm{c})\bigr)\bigr] > 0. \tag{2}$$

The sign of this derivative suggests that reductions in search effort resulting from negative neighborhood effects should reduce employment probabilities for any given t, thereby increasing the duration of unemployment. Neighborhood effects may also alter job offer arrival rates directly if neighbors are a source of information about potential jobs. In this case, reductions in α resulting from living in a segregated environment are associated with a lower probability of accepting employment, provided that α does not affect reservation wages. Job offer arrival rates may also be lower for black residents of segregated neighborhoods if suburban employers discriminate based on the residential location of prospective employees.

For a given level of search effort, residential segregation may also alter reservation wages through changes in job search costs and commuting costs. Job search costs may be higher for minority workers residing in segregated neighborhoods if segregated neighborhoods are spatially separated from the location of job openings. Given an acceptable wage offer, spatially distant jobs also entail higher commuting costs. Since job search costs and commuting costs are likely to be lower for those who own automobiles than for those who rely on public transit to search for and commute to employment [24], automobile ownership status likely mitigates the impact of spatial barriers on unemployment durations.

As shown by Rogers [16], changes in job search costs and commuting costs each have offsetting impacts on employment probabilities. Increased job search costs lower reservation wages required to accept employment, thereby increasing the probability of employment. Commuting costs, on the other hand, increase reservation wages and reduce the probability of accepting a given wage offer at time t. The spatial impacts of segregation on the probability of leaving unemployment at time t, then, depend on the interaction among these effects. *Ceteris paribus*, an increase in segregation which reduces offer arrival rates will reduce the probability of leaving unemployment at time t, provided that the impact of increased commuting costs is at least as large as the impact of increased job search costs. Due to the interaction among these four effects (search effort, job offer arrival rates, search costs, and commuting costs), it is difficult to determine theoretically how segregation may impact unemployment transition probabilities, thus pointing to the need for empirical analysis of these effects.

Empirical evidence on the importance of spatial barriers and neighborhood effects on unemployment probabilities has produced mixed results. Several empirical studies provide support for some version of the spatial mismatch hypothesis linking poor job accessibility for black workers to relatively lower employment probabilities (see the reviews by Ihlanfeldt and Sjoquist [7] and Preston and McLafferty [13]). Others minimize the importance of residential location in influencing labor market outcomes (see, for example, Ellwood [2]) or point to the prevailing importance of auto ownership in determining low-income workers' job accessibility and subsequently employment status (see, for example, Ong [12], Kawabata [9], Sanchez et al. [18]). Similarly inconclusive findings regarding the importance of neighborhood effects have been obtained. Jencks and Mayer [8] review early literature and find little evidence to support the existence of significant neighborhood effects, while a review of more recent evidence finds support for the neighborhood effects argument [3]. Of those studies reviewed by both authors, few examine the impacts of neighborhood effects on adult labor market outcomes [11].

Of those studies that have examined the impact of residential location on racial differences in unemployment within US metropolitan areas, only two focus exclusively on unemployment durations. Of these, neither examines the neighborhood effects described by Wilson [25]. Holzer, Ihlanfeldt, and Sjoquist [6] rely on data from the 1981 and 1982 National Longitudinal Survey of Youth (NLSY) to estimate models explaining the log of unemployment duration for a sample of black and white youths. The authors find that automobile ownership shortens unemployment durations for a pooled sample of all households, and a measure of job decentralization is associated with increased unemployment durations. Transportation mode and job decentralization explain approximately one third of the racial difference in unemployment durations. Furthermore, in racially-stratified models, the authors find that job decentralization has a larger impact on the duration of unemployment for blacks than for whites. Their examination of unemployment durations does not control for measures of worker characteristics other than transportation mode and does not control for measures of neighborhood effects. Furthermore, the models are estimated using ordinary least squares (OLS), which does not take unique features of the distribution of duration times into account, including non-negativity, non-linearities with respect to time, and possible censoring of duration times.

Rogers [16] relies on unemployment insurance claim data for the Pittsburgh, Pennsylvania, metropolitan area in 1980 through 1986 to examine the impact of job accessibility on unemployment durations. In the estimated equations, she employs measures of job accessibility and local unemployment rates along with other human capital controls. Although maximum likelihood procedures suitable for explaining unemployment durations are employed, the sensitivity of the results to the parameterization of the probability density function of duration times is not examined. Job accessibility is measured using two different employment-based indices that are each weighted by commuting times between the location of a household's residence and the location of jobs. Although access to employment growth is found to reduce unemployment duration, the results are sensitive to the functional form of the accessibility measure. Her MSA-level study does not report results from racially stratified models and does not include controls for the neighborhood effects mentioned by Wilson [25].

Our research builds upon these two earlier studies by simultaneously controlling for measures of neighborhood peer effects, job accessibility, vehicle ownership status, along with a variety of other household and worker characteristics affecting unemployment durations to determine the marginal impact of each on unemployment durations. We also employ more precise measures of job accessibility than other multiple-MSA studies examining the link between residential location and racial differences in employment outcomes.

3. Empirical strategy and data description

To disentangle the relative contribution of job accessibility, neighborhood peer effects, and vehicle ownership status towards racial differences in unemployment durations, we rely on continuous time duration models to examine the transition to employment among those who have recently been dismissed from a job. In the empirical analysis, we employ a unique version of the Panel Study of Income Dynamics (PSID) that provides information on the census tract and zip code of residence for each individual in the sample. Geocodes linking individuals to their residential location were acquired from the Institute for Social Research (ISR) at the University of Michigan. The final sample consists of approximately 1000 workers (household heads and spouses of household heads) who identify their primary race as "white" or "black,"[1] who were unemployed during the period 1990 through 1992, and who resided in metropolitan statistical areas (MSAs) throughout the period under investigation. From the date of initial employment termination, we follow individuals until their date of reemployment and record the length of their initial unemployment duration in months. Since workers who identified themselves as being out of the labor force were omitted from the analysis, we focus only on those who were actively seeking employment. Characteristics of the individual's residential location were obtained by matching workers to the residential location where the worker resided at the time of job loss.[2]

Using this database, we examine several quantities of interest related to the distribution of unemployment durations. If we define t as a random variable which represents a given individual's unemployment spell, the probability density function of unemployment durations is then defined as $f(t)$. The survival function, $S(t)$, gives the probability that person i is unemployed until at least time t. The hazard rate, $\lambda(t)$, then gives the conditional proba-

[1] An earlier analysis also examined the unemployment durations of Hispanic workers. Since the majority of Hispanics in this sample were included as part of a special Latino supplement to the PSID added in 1990 and were not included in the 1989 wave of the PSID, their vehicle ownership status and wealth status could not be determined. When those included in the Latino supplement were dropped from the sample, there were too few Hispanic workers to obtain reliable estimates from racially stratified models.

[2] The data used in this analysis are derived from Sensitive Data Files of the Panel Study of Income Dynamics, obtained under special contractual arrangements designed to protect the anonymity of respondents. These data are not available from the authors. Persons interested in obtaining PSID Sensitive Data Files should contact the ISR through the Internet at PSIDHelp@isr.umich.edu. Census tract information is matched to each PSID worker using geocode match files, which identify the census tract and zip code of residence for each worker in the sample. Census tract-level variables are obtained from the Neighborhood Change Database, distributed by Geolytics, Inc.

bility of leaving an unemployment state, given unemployment until time t. The hazard rate is constructed as follows: $f(t)/S(t)$. In the models that follow, we define $f(t)$ and $S(t)$ to be a function of a vector of independent variables (X) and parameters (β) and choose β to maximize the log of the following likelihood function, $L(\beta)$:

$$L(\beta) = \prod_{i=1}^{n} f(t_i, X\beta)^{\delta_i} \prod_{i=1}^{n} \bigl[1 - S(t_i, X\beta)\bigr]^{1-\delta_i}. \tag{3}$$

Since several workers in the sample were unemployed at the beginning of the observation period, the likelihood function is modified to account for the left-censored unemployment durations for these workers. If we define a censoring indicator, δ_i, which is equal to 1 if the household's date of entry into unemployment is observed and 0 otherwise, the likelihood can be constructed as the product of the density function for workers with uncensored spells $(\prod_{i=1}^{n} f(t_i, X\beta))$ and one minus the survival function for workers with censored spells $(\prod_{i=1}^{n}[1 - S(t_i, X\beta)])$. Several parameterizations of (3) are examined to determine which provides the best fit. All models are estimated using the S-PLUS statistical software package.

To explain unemployment durations for the ith worker, we include race, gender, and other human capital variables that affect worker productivity and earnings potential (age, health status, household income prior to unemployment, total family wealth prior to unemployment, education, number of children, age of youngest child, marital status, vehicle ownership status, industry controls, and year of employment termination). Each of these variables, with the exception of wealth and vehicle ownership status, is based on measured values as of 1990, the beginning of the sample period.

Several variables deserve some elaboration. Health status is measured using a dummy variable equal to one if the PSID respondent reported having "good" health or better and zero otherwise. We also include several dummy variables indicating the worker's industry of employment just prior to employment termination. These industry controls were created by collapsing the worker's 1970 3-digit SIC industry code into one of four major categories: retail, manufacturing, service, and construction, with the omitted category including all workers employed in all other industries (agriculture, mining, transportation, finance, insurance, real estate, and public administration). Finally, we include dummy variables indicating the year in which the unemployment spell began to capture temporal macroeconomic factors affecting unemployment durations.

Two variables were extracted from the 1989 wealth supplement of the PSID, which provides data on the dollar value of various types of assets held by each household in the sample. The first is an indicator of vehicle ownership status. All households who reported a positive dollar value of vehicle assets in 1989 were identified as auto owners. Family wealth is measured as the dollar value of all non-housing assets owned minus the debt owed on those assets by the family in 1989.

We also include several variables measured at the census tract level for each worker in the sample. Two direct measures of neighborhood peer effects are included (percent of census tract residents who are below the poverty line and percent of census tract residents with high school degrees). "Neighborhood percent non-Hispanic white" captures other neighborhood-level effects correlated with racial segregation that are not captured by the two direct measures of neighborhood effects. For example, racially prejudiced employers may discriminate based on the racial composition of one's residential neighborhood. Finally, we include a dummy variable equal to one if the worker resides in the central city and zero otherwise. Our objective is not to estimate a parsimonious structural model of the spatial factors contributing to unemployment but rather to examine how key indicators of neighborhood-level conditions impinge upon one's employment transition.

All models also include a measure of job accessibility, defined as the share of total MSA jobs located in each worker's zip code of residence.[3] Although this measure incorporates less information on the spatial distribution of jobs than gravity-based measures employed by others, such as Kawabata [9], Raphael [14], Sanchez et al. [18] and Shen [20,21], our sample includes workers across 158 different metropolitan areas. The large number of metropolitan areas in the sample makes the computation of gravity-based measures computationally difficult. As such, we opt for the simpler, "zip code share of MSA jobs." Despite the crudeness of this measure, our measure improves upon other similar multi-MSA studies, such as Holzer et al. [6] and Ross [17], who measure job accessibility in terms of the total share of metropolitan jobs located in the central city, interacted with an indicator of the worker's residential location inside the central city. In Appendix B, we compare the estimates from models utilizing our chosen measure of accessibility to other similar measures utilized in the literature, such as a total job measure, a job/worker measure, and commuting-based measures.

All census tract and zip code-level variables are based on measured values as of the date of the worker's employment termination. Thus, if the worker became unemployed in 1992, for example, that worker's census tract (or zip code) characteristics would be measured by matching the appropriate census tract (or zip code) geo-code from the 1992 wave of the PSID to the appropriate census tract (or zip code) identifier for that worker. We adopt this procedure to ensure that workers who moved between the initial sample period (1990) and the date of employment termination are assigned the location characteristics of their most recent residential location just prior to employment termination.

None of the variables utilized in the analysis vary with time, including the residential location characteristics. We do not examine models which consider time-varying covariates for several reasons. First, the assumption of time-constant covariates simplifies the construction of the database and the estimation of the parameters of the log likelihood function. Furthermore, since most unemployment durations in the sample (89 percent) were less than one year in duration, allowing for time-varying covariates would only provide additional information for a small minority of workers in the sample. Given that PSID workers are resurveyed on an annual basis only, it is not clear that allowing for time-varying covariates would necessarily reduce errors in measurement in this particular case.

[3] Zip code job information is obtained from the 1987 Censuses of Manufactures, Retail, and Service Industries and matched to each PSID worker using 1990 zip code identifiers. Since the 1987 and 1992 Censuses of Manufactures, Retail, and Service Industries report job totals for manufacturing, retail, and service industries only, our job total measure refers to jobs within these categories only. Furthermore, since the Census of Manufactures reports establishment-level data rather than employment-level data, we impute total manufacturing jobs using the procedure suggested by Raphael and Stoll [15].

To determine how neighborhood characteristics affect the racial differential in unemployment durations, we examine the sign and significance of the race coefficient in models with and without neighborhood characteristics to determine how neighborhoods affect the impact of race on unemployment durations. We also examine racially-stratified models to determine if the impact of residential location characteristics on unemployment durations varies by race.

4. Results

4.1. Descriptive statistics

Table 1 displays descriptive statistics for the full sample and for each racial group separately. As this table suggests, the average worker in the sample is unemployed for approximately 5 months, is approximately 35 years of age, is likely to own an automobile, is likely to be married, and is more likely to have earned a high school degree than a college degree. The distribution of workers across industry types is relatively uniform across all industries except the construction industry, with the preponderance of workers employed in service and other industries. During the period under investigation, the majority of workers began their unemployment spells in either 1990 or 1991.

On average, white workers experience shorter unemployment durations than the black workers in the sample. Table 1 also points to significant racial differences in several worker characteristics, including family income, family wealth, health status of the worker, automobile ownership status, marital status, education level, and industry composition. Note, in particular, that while 90 percent of whites own automobiles, only 53 percent of black workers own autos. Furthermore, whites live in neighborhoods that are less likely to be located in the central city, where the majority of neighbors are white, where poverty rates are low, where high school graduation rates are high, and where jobs are relatively more accessible relative to the surrounding metropolitan area.

When we examine differences in the job accessibility measure by race and by central city status (results not shown for brevity), we found that while blacks living in the suburbs have job accessibility measures that are comparable to whites (0.034 for white workers versus 0.027 for black workers), blacks living in the central city have much poorer levels of job accessibility than do whites living in the central city (0.087 for white workers versus 0.032 for black workers). This has important implications. First, it suggests that measures of job decentralization (share of the MSA's jobs located in the central city) employed by previous studies (see, for example, [6]) may obscure differences in accessibility *within* central cities and suburban areas. Second, these results are consistent with and extend the results of Shen [20,21] who finds that central city residents actually have higher measures of job accessibility when job accessibility measures incorporate differences in commuting mode and jobs resulting from turnover versus new job creation. Our findings suggest that the central city jobs resulting from job turnover may be more accessible to white workers than to black workers. Our measure of accessibility does not incorporate travel distance or cost, however, and does not differentiate among different modes of commuting.

Table 1
Descriptive statistics

Variable	Full sample $N = 1098$		White $N = 558$		Black $N = 540$	
	Mean	Std. dev.	Mean	Std. dev.	Mean	Std. dev.
Dependent variable						
Unemployment duration (months)	4.830	4.871	4.081	4.215	5.604	5.362
Uncensored	0.788	0.409	0.842	0.365	0.731	0.444
Independent variables: individual & HH characteristics						
Female	0.276	0.447	0.357	0.479	0.193	0.395
Family income, 1989 (1000$s)	29.418	24.805	39.685	26.490	18.810	17.473
Family wealth, 1989 (1000$s)	24.103	126.235	36.802	160.414	10.981	74.170
Age	35.384	10.439	36.466	10.633	34.267	10.124
Good health	0.249	0.432	0.294	0.456	0.202	0.402
Own automobile	0.715	0.452	0.898	0.303	0.526	0.500
Married	0.587	0.493	0.746	0.436	0.424	0.495
College graduate	0.109	0.312	0.183	0.387	0.033	0.180
High school graduate	0.777	0.417	0.867	0.339	0.683	0.466
# Children in family	1.172	1.336	1.023	1.175	1.326	1.470
Age of youngest child in family	3.430	4.669	3.495	4.890	3.363	4.433
Independent variables: industry controls						
Retail	0.200	0.400	0.204	0.404	0.196	0.398
Manufacturing	0.175	0.380	0.224	0.417	0.124	0.330
Service	0.267	0.443	0.267	0.443	0.267	0.443
Construction	0.087	0.283	0.093	0.291	0.081	0.274
Other	0.270	0.444	0.211	0.409	0.331	0.471
Independent variables: duration dummy variables						
Unemployment spell began 1990	0.477	0.500	0.430	0.496	0.526	0.500
Unemployment spell began 1991	0.333	0.472	0.349	0.477	0.317	0.466
Unemployment spell began 1992	0.189	0.392	0.220	0.415	0.157	0.365
Independent variables: location controls						
Central city	0.548	0.498	0.328	0.470	0.776	0.417
% White non-Hispanic	0.547	0.380	0.826	0.201	0.259	0.299
% Poor	0.191	0.165	0.100	0.093	0.285	0.170
% High school education	0.704	0.161	0.781	0.130	0.625	0.152
Zip code share of total MSA jobs	0.041	0.086	0.052	0.108	0.031	0.051

4.2. Initial model diagnostics

We initially examined several different parameterizations of the likelihood function shown in Eq. (3). Since the distribution of duration times for the workers in the sample is non-negative, possibly increasing or decreasing with time, and left-censored, ordinary least squares estimation is generally not appropriate. We consider a general class of models known as “accelerated failure time models,” which specify the distribution of duration times as follows:

$$Y = X'\beta + \sigma\varepsilon, \tag{4}$$

where Y is the log of unemployment duration, X is a vector of covariates, β is an estimated parameter vector, ε is an error term, and σ is an estimated scale parameter. Written in this manner, different parametric models can be derived by imposing different distributional assumptions on ε and by allowing σ to assume different values. If ε is distributed as extreme value and $\sigma = 1$, (4) reduces to an exponential model. Similarly, if $\sigma = 0.5$ and ε is distributed as extreme value, (4) reduces to the Rayleigh distribution. The Weibull distribution is a more general distribution that allows for positive or negative duration dependence, with the degree of duration dependence captured by the estimated parameter, σ. When $1/\sigma < 1$, the baseline hazard function monotonically decreases with time (negative duration dependence), and when $1/\sigma > 1$, the baseline hazard function monotonically increases with time (positive duration dependence).

Two other distributions for duration times were also examined: (1) the normal distribution, which is equivalent to the standard Tobit model when censoring is taken into account, and (2) the logistic distribution. The logistic distribution is examined primarily to provide a basis for comparison with Rogers [16]. Other more complex models such as those based on the generalized F distribution were also examined but were ultimately abandoned due to convergence problems.

To determine the best fitting functional form, we examined Akaike's information criterion (AIC), calculated as: $-2 * (\log \text{ likelihood}) + (p * k)$, where $p =$ number of parameters estimated and $k =$ constant chosen by the analyst. Although there is some disagreement over the appropriate choice of k, most recommend $k = 3$ for general model comparisons [23]. Other than graphical techniques, the AIC provides one of the only diagnostic tools that can be used to compare non-nested parametric models, such as the ones considered above. As shown in Table 2, the Weibull model has the lowest AIC and the log likelihood that is closest to zero. Furthermore, since the Rayleigh and exponential distributions are both nested within the Weibull, we can rely on likelihood ratio tests to evaluate the appropriateness of the restrictions imposed on the scale parameter. Both tests soundly favor the Weibull specification over the exponential and Rayleigh specification. Based on these results, we choose the Weibull functional form for the remainder of our parametric results. As shown in Appendix A, alternative parameterizations of the log likelihood function produce results that are highly comparable to those obtained from the Weibull model.

Table 2
Unemployment duration initial model comparisons[a]

Parameterization	Scale	Log likelihood	AIC
Weibull	0.90	−2227.90	4530.80
	(0.03)		
Exponential	1.00	−2236.20	4544.40
Rayleigh	0.50	−2619.90	5311.80
Logistic	2.16	−2528.20	5131.40
	(0.03)		
Normal	4.44	−2618.40	5311.80
	(0.02)		

[a] Standard errors for scale parameter estimate shown in parentheses.

4.3. Full sample results

We now turn to an examination of regression models that explain the determinants of unemployment durations for individual workers in the sample. Results from four different models relying on the Weibull distribution are displayed in Table 3. Model 1 is a baseline model with controls for race only. Model 2 controls for observable household and worker characteristics; Model 3 controls for household and worker characteristics plus a measure of job accessibility; and Model 4 controls for the full set of worker, household, job accessibility, and neighborhood characteristics. The coefficient on the race variable in each of these equations gives an estimate of the racial differential in unemployment times, conditioning on various determinants of unemployment durations.

Although most traditional pseudo-R^2-type measures such as McFadden's R^2 have been shown to be inadequate for evaluating the goodness-of-fit of parametric survival models which exhibit censoring [19], Kiefer [10] suggests, at a minimum, examining likelihood ratio tests of the significance of the regression model versus an intercept-only model. All such tests soundly reject the intercept-only model. The estimated scale parameter (σ) of 0.901 in the full model provides evidence of positive duration dependence. This suggests that the average, or "baseline," household is slightly more likely to transition out of unemployment with longer unemployment spells.

Since the partial derivatives of unemployment duration with respect to each independent variable are non-linear functions of the βs, the Xs, and σ, the coefficients from the models cannot be directly interpreted as marginal effects. The sign of the estimated coefficients can be interpreted in terms of the accelerated failure time model, however. Here, positive coefficients indicate that the variable contributes to longer unemployment durations, while negative coefficients indicate that the variable contributes to shorter unemployment durations. The exponential of the estimated coefficients can also be roughly interpreted as the percentage change in unemployment duration resulting from a unit increase in the independent variable.

Regarding the impact of covariates on unemployment durations, we find that in the full model, unemployment durations are reduced with improved health status, vehicle ownership status, increased numbers of children and younger children, employment in service industries, earlier employment termination date, and increased job accessibility. The importance of presence of children is consistent with previous research examining the determinants of employment status [18]. Controlling for the full set of worker characteristics, we find that a worker's previous income does not have a statistically significant impact on his or her eventual duration of unemployment, a finding that is consistent with Rogers [16]. The impact of wealth on unemployment durations is also statistically insignificant. Our findings regarding the impact of income and wealth also hold in models which consider income and wealth separately.

Although, in general, vehicle ownership status is statistically significant across all model specifications, differences in the magnitude of the vehicle ownership status coefficient are particularly interesting. Although owning a vehicle contributes to a shorter unemployment duration overall, the pattern of results suggests that residential location characteristics mitigate the impact of vehicle ownership status on unemployment durations. In the model with the most extensive set of residential location controls, the vehicle ownership status coef-

Table 3
Weibull unemployment duration models, full sample ($N = 1098$)

Variable	Model 1		Model 2		Model 3		Model 4	
	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error
Race of worker								
Constant	1.296	0.043***	1.032	0.166***	1.158	0.168***	1.289	0.312***
Black	0.186	0.059***	0.173	0.069**	0.140	0.069**	0.067	0.091
Individual & HH characteristics								
Female	–	–	0.082	0.080	0.071	0.080	0.075	0.080
Family income, 1989 (1000$s)	–	–	0.001	0.002	3.2E-04	0.002	7.0E-05	0.002
Family wealth, 1989 (1000$s)			−1.9E-04	2.6E-04	−2.0E-04	2.6E-04	−2.0E-04	2.5E-04
Age	–	–	0.005	0.003	0.004	0.003	0.004	0.003
Good health	–	–	−0.163	0.071**	−0.171	0.070**	−0.166	0.070**
Own automobile	–	–	−0.151	0.077**	−0.146	0.076*	−0.136	0.077*
Married	–	–	−0.082	0.083	−0.062	0.082	−0.051	0.082
College graduate	–	–	0.080	0.097	0.069	0.096	0.072	0.097
High school graduate	–	–	0.015	0.075	0.002	0.075	−0.002	0.075
# Children in family	–	–	−0.055	0.025**	−0.051	0.024**	−0.055	0.024**
Age of youngest child in family	–	–	0.019	0.007***	0.018	0.007***	0.018	0.007***
Industry controls								
Retail	–	–	−0.075	0.086	−0.090	0.085	−0.080	0.085
Manufacturing	–	–	−0.158	0.090*	−0.125	0.091	−0.104	0.091
Service	–	–	−0.208	0.080***	−0.197	0.079**	−0.177	0.079**
Construction	–	–	−0.103	0.116	−0.107	0.115	−0.098	0.117
Duration dummy variables								
Unemployment spell began 1991	–	–	0.545	0.066***	0.535	0.065***	0.543	0.065***
Unemployment spell began 1992	–	–	0.574	0.079***	0.571	0.078***	0.562	0.078***

(*continued on next page*)

Table 3 (*continued*)

Variable	Model 1		Model 2		Model 3		Model 4	
	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error
Location controls								
Central city	–	–	–	–	–	–	−0.103	0.074
% White non-Hispanic	–	–	–	–	–	–	−0.208	0.150
% Poor	–	–	–	–	–	–	0.104	0.336
% High school education	–	–	–	–	–	–	0.049	0.314
Zip code share of total MSA jobs	–	–	–	–	−1.311	0.335***	−1.142	0.347***
Scale	0.963	0.025	0.914	0.025***	0.905	0.025***	0.901	0.025***
Log-likelihood	−2297.6		−2236.500		−2229.900		−2227.900	
χ^{2a}	9.800	***	132.000	***	145.200	***	149.200	***

[a] LR test of model significance.
* $p < 0.10$.
** $p < 0.05$.
*** $p < 0.01$.

ficient is roughly 10 percent smaller in magnitude than in the model that does not control for location characteristics. Furthermore, vehicle ownership status is significant only at the 0.10 level once location characteristics are controlled for. This result suggests that auto ownership is likely correlated with and mitigated by the residential location characteristics of workers.

The importance of job accessibility is readily apparent. In the full model with the most extensive set of controls, job accessibility is significant at the 0.01 level. Furthermore, this effect is nontrivial in magnitude. A one standard deviation increase (0.086) in local shares of MSA jobs equates with a roughly 9 percent reduction in unemployment durations. By comparison, automobile ownership status reduces unemployment durations by approximately 13 percent, a reduction that is roughly equivalent to the impact of a 12 percentage point increase in job accessibility.

No other residential location characteristic is statistically significant in the full model. Obviously, this partly reflects the high inter-correlations among the neighborhood characteristics in the model. Bivariate correlations among the chosen neighborhood characteristics are all greater than 0.40 in magnitude. To determine if neighborhood characteristics were significant individually, we also ran models entering each neighborhood characteristic separately in models with and without controls for job accessibility. In these models, "neighborhood percent non-Hispanic white" was the only statistically significant neighborhood characteristic ($p < 0.10$). This variable is no longer significant once we control for job accessibility, however.

Differences in the race indicator variables across the three models provide evidence on the conditional impact of residential location on racial differences in unemployment durations. In the model without covariates, black workers are unemployed for approximately 20 percent longer than white workers. The black-white differential declines by 7 percent once we control for worker and household characteristics, 25 percent when we control for job accessibility, and 64 percent in models with the full set of residential location controls. Furthermore, in the full model, the racial differential in unemployment durations is no longer statistically different from zero.

To determine the sensitivity of these results to the chosen measure of job accessibility, we re-estimated Model 4 using several other zip code-based measures of job accessibility, including zip code shares of total MSA manufacturing jobs, zip code shares of total MSA retail jobs, zip code shares of total service jobs, total jobs per zip code, and ratio of jobs to resident workers for each zip code. We also examined several census tract commuting-based measures and a measure of public transit utilization within each worker's surrounding census tract. Controlling for other residential location characteristics, only the zip code-based measures are statistically significant at the 0.10 level. These estimates are provided in Appendix B.[4]

[4] We also examined models for the subset of workers whose unemployment durations were not left-censored. Although the sample size was substantially smaller in these regressions, the results were qualitatively similar to those reported using the full sample of censored and uncensored observations. Additionally, we examined models for central city workers only to determine if the spatial delineation of zip code boundaries possibly affects our estimates. Again, the results were qualitatively similar to those reported above, so retain the full sample of central city and non-central city cases.

4.4. Unemployment durations by race

In this section, we examine racially-stratified models of unemployment duration to determine if residential location characteristics affect blacks and whites differently. These results are displayed in Table 4.

These results largely confirm the results from earlier regressions with a few exceptions. In general, we find evidence of racial heterogeneity in the impact of covariates on unemployment durations. Among whites, age, health status, number of children, age of

Table 4
Racially-stratified unemployment duration models

Variable	Black workers ($N = 540$)		White workers ($N = 558$)	
	Coef.	St. error	Coef.	St. error
Individual & HH characteristics				
Constant	1.043	0.449**	1.699	0.486**
Female	0.507	0.134**	−0.059	0.096
Family income, 1989 (1000$s)	0.005	0.003	−0.001	0.002
Family wealth, 1989 (1000$s)	−2.1E-04	4.8E-04	−2.9E-04	3.4E-04
Age	−0.002	0.004	0.007	0.004*
Good health	0.061	0.109	−0.252	0.090***
Own automobile	−0.211	0.099**	−0.068	0.132
Married	−0.317	0.120***	0.073	0.115
College graduate	−0.405	0.227*	0.137	0.104
High school graduate	0.147	0.097	−0.195	0.123
# Children in family	−0.025	0.031	−0.116	0.041***
Age of youngest child in family	0.019	0.010*	0.023	0.009***
Industry controls				
Retail	0.021	0.120	−0.188	0.121
Manufacturing	−0.139	0.142	−0.157	0.118
Service	−0.048	0.110	−0.264	0.115**
Construction	0.037	0.172	−0.178	0.158
Duration dummy variables				
Unemployment spell began 1991	0.707	0.093***	0.402	0.088***
Unemployment spell began 1992	0.624	0.117***	0.575	0.102***
Location controls				
Central city	−0.046	0.110	−0.143	0.096
% White non-Hispanic	−0.078	0.193	−0.337	0.254
% Poor	0.440	0.420	−0.671	0.724
% High school education	0.074	0.470	−0.005	0.438
Zip code share of total MSA jobs	−2.168	0.971**	−1.025	0.377***
Scale	0.885	0.037***	0.859	0.034***
Log-likelihood	−1070.400		−1131.400	
$\chi^{2\text{a}}$	102.000	***	89.000	***

[a] LR test of model significance.
* $p < 0.10$.
** $p < 0.05$.
*** $p < 0.01$.

youngest child, employment in service industries, year of employment termination, and job accessibility are all significant predictors of unemployment durations. Among black workers, gender, automobile ownership status, marital status, educational attainment, age of youngest child, year of employment termination, and job accessibility are statistically significant determinants of unemployment duration.

Several of these findings provide information concerning the impact of job accessibility and neighborhood characteristics on racial differences in unemployment durations. First note that automobile ownership status is statistically significant for black workers only. If African American workers reside in locations that are more distant from employment locations, or rely more heavily on public transportation to reach employment opportunities, automobile ownership would be expected to have the largest impact on employment for these workers, who would stand to benefit most from automobility. This finding is further supported by the relatively larger impact of job accessibility among black workers. The impact of job accessibility is more than twice as large for black workers than for white workers, a finding that is consistent with Holzer et al. [6].

Other than job accessibility, no other residential location characteristic is statistically significant. To determine if this finding is affected by high inter-correlations among the independent variables in the model, we examined additional models which control for each neighborhood characteristic separately, with and without controls for job accessibility. Estimates from these models are summarized in Table 5. All models include the full set of controls shown in Model 2, Table 3, with the addition of different combinations of residential location characteristics. Model 1, for example, includes controls for central city status and the zip code share of total MSA jobs. Since all 8 models are estimated for separate samples of black and white workers, Table 5 displays the results from 16 different regressions.

When entered separately, all neighborhood characteristics except central city status affect unemployment durations for blacks. Unemployment durations for blacks are reduced with increased exposure to non-Hispanic whites, lower poverty rates, and higher exposure to high-school educated neighbors. With controls for job accessibility, only neighborhood poverty rates significantly affect unemployment durations.

Among whites, job accessibility is the only location control affecting unemployment durations. Interestingly, however, residence inside the central city reduces unemployment durations for white workers. There are several possible explanations for this unexpected finding. First, white workers may have been more highly represented in higher-wage professional occupations prior to employment termination. These jobs are still heavily centralized in the central city of most metropolitan areas. Second, as pointed out in the discussion of descriptive statistics above, whites living in the central city have much higher levels of job accessibility than do blacks living in the central city. Third, blacks living in the central city may rely on public transit to conduct job search, whereas whites may rely primarily on automobiles. This finding clearly deserves further exploration in future studies.

The estimated racial differences in the importance of neighborhood effects paint a picture that is similar to that suggested by the job accessibility estimates. As suggested by Galster et al. [4], negative neighborhood effects may not significantly affect opportunity structures until such effects have reached critical threshold levels. If these critical thresholds are never reached within majority-white neighborhoods, then it is reasonable that

Table 5
Coefficients from racially-stratified models with residential location controls entered separately[a]

Sample	Residential location controls	Model							
		1 Coef. (Std. error)	2 Coef. (Std. error)	3 Coef. (Std. error)	4 Coef. (Std. error)	5 Coef. (Std. error)	6 Coef. (Std. error)	7 Coef. (Std. error)	8 Coef. (Std. error)
Black workers									
	Central city	0.033				0.007			
		(0.102)				(0.101)			
	% White non-Hispanic		-0.185				-0.257^{*}		
			(0.149)				(0.144)		
	% Poor			0.418^{*}				0.461^{*}	
				(0.251)				(0.252)	
	% High school education				-0.365				-0.469^{*}
					(0.279)				(0.277)
	Zip code share of total MSA jobs	-2.434^{***}	-2.128^{**}	-2.300^{**}	-2.199^{**}				
		(0.920)	(0.948)	(0.923)	(0.935)				
White workers									
	Central city	-0.132				-0.181^{**}			
		(0.088)				(0.087)			
	% White non-Hispanic		-0.055				-0.103		
			(0.198)				(0.200)		
	% Poor			-0.429				-0.432	
				(0.458)				(0.461)	
	% High school education				0.132				0.091
					(0.326)				(0.330)
	Zip code share of total MSA jobs	-1.091^{***}	-1.198^{***}	-1.211^{***}	-1.215^{***}				
		(0.370)	(0.362)	(0.362)	(0.361)				

[a] All models include full set of controls shown in Model 2, Table 3, with the addition of each residential location control above.
* $p < 0.10$.
** $p < 0.05$.
*** $p < 0.01$.

changes in neighborhood characteristics would have little impact on the unemployment transitions of white workers. Black workers, on the other hand, may be more likely to reside in neighborhoods where negative peer effects have reached critical thresholds. Since black workers stand to benefit most from proximity to non-poor, educated neighbors, altering the residential locations of black workers can be expected to have the largest impact on overall unemployment durations.

We qualify our findings by pointing out that our results may be biased by the endogeneity of residential location for unemployed workers. Workers who choose to live in high-poverty neighborhoods, for example, may have a different labor market orientation than those who choose to live in low-poverty neighborhoods. Therefore, the effect of neighborhood characteristics on unemployment duration may jointly capture neighborhood effects and the unobservable characteristics of workers. Although we partially address this issue by focusing on the residential location of workers at the time of employment termination—based on the assumption that such neighborhoods were chosen prior to employment termination—unobservable time-invariant worker characteristics affecting previous location decisions will still be correlated with future unemployment outcomes. Further analysis is required to determine the sensitivity of our results to possible self-selection bias.

5. Conclusion

This paper examined the connection between residential location, automobile ownership status, and racial differences in unemployment durations. Using worker-level data from the Panel Study of Income Dynamics, linked to data describing the residential locations of workers in the sample, we find that residential segregation affects racial differences in unemployment durations by exacerbating racial differences in job accessibility and neighborhood effects. For the average worker in the sample, we find that residential location characteristics tend to mitigate the impact of auto ownership on unemployment durations. Furthermore, the racial differential in unemployment durations disappears once we control for residential location characteristics.

In racially-stratified models, we find that the impact of residential location and vehicle ownership status varies considerably across racial groups. For African Americans, vehicle ownership status and increased job accessibility both significantly reduce unemployment durations. Among whites, only job accessibility reduces unemployment duration. Comparing the unemployment durations of whites and blacks, we find that the impact of job accessibility is more than twice as large for blacks as it is for whites. Furthermore, neighborhood peer effects are shown to be statistically significant determinants of unemployment durations for black workers only.

These findings provide new evidence on suggested policy alternatives for reducing racial gaps in unemployment. Our evidence suggests that while automobile subsidies may be an effective means of increasing the job accessibility of African Americans, the impact of these subsidies will tend to be mitigated by neighborhood conditions if African Americans continue to reside in neighborhoods with poor job accessibility, few positive role models, and few networks to new job sources. Alternatively, automobile subsidies are arguably an effective means of increasing the job accessibility of African Americans given the persistence of racial segregation which isolates low income black workers to high poverty locations.

Policies to improve housing opportunities for minorities in job-rich areas are likely to be successful, especially if such initiatives focus on alleviating housing market discrimination. Given our finding that the average worker's previous income does not contribute to his or her eventual unemployment duration, controlling for other correlates of worker productivity, policies which focus solely on affordability are less likely to be as successful as policies designed to remove discriminatory constraints on minority housing choice. Finally, economic development strategies focusing on education within minority communities are likely to have a significant impact on African American unemployment durations, both directly through increases in human capital and indirectly through network effects.

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Appendix A

See Table A.1.

Appendix B

Table A.2 displays the coefficients from several models utilizing different measures of job accessibility. Job accessibility measures include:

- zip code share of total MSA jobs;
- zip code share of total MSA manufacturing jobs;
- zip code share of total MSA retail jobs;
- zip code share of total MSA service jobs;
- total jobs in zip code;
- total zip code jobs/total workers residing in zip code;
- % of census tract workers with commute times $>$ 25 minutes;
- % of census tract workers who work within the surrounding county;
- % of census tract workers who travel to work on public transit.

We also examined zip code based measures utilizing 1992 job estimates instead of 1987 estimates. The results were very similar to those shown above, so we opt to employ the lagged job measure. We also examined measures of changes in jobs between 1987 and

Table A.1
Unemployment duration models, alternative parameterizations ($N = 1098$)

Variables	Parameterization									
	Weibull		Exponential		Rayleigh		Logistic		Normal	
	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error
Race of worker										
Constant	1.289	0.312***	1.232	0.344***	1.579	0.178***	3.332	1.284***	3.391	1.560**
Black	0.067	0.091	0.068	0.100	0.073	0.052	0.341	0.364	0.256	0.448
Individual & HH characteristics										
Female	0.075	0.080	0.093	0.088	−0.043	0.047	0.745	0.318**	0.535	0.394
Family income, 1989 (1000$s)	7.0E-05	0.002	2.0E-04	0.002	−0.001	0.001	0.002	0.007	0.003	0.008
Family wealth, 1989 (1000$s)	−2.0E-04	2.5E-04	−2.0E-04	2.8E-04	−2.6E-04	1.5E-04*	−3.2E-04	0.001	−0.001	0.001
Age	0.004	0.003	0.004	0.003	0.005	0.002***	0.005	0.012	0.008	0.015
Good health	−0.166	0.070**	−0.166	0.077**	−0.168	0.041***	−0.432	0.281	−0.664	0.347*
Own automobile	−0.136	0.077*	−0.134	0.085	−0.140	0.043***	−0.473	0.330	−0.468	0.396
Married	−0.051	0.082	−0.067	0.091	0.068	0.047	−0.596	0.342*	−0.487	0.419
College graduate	0.072	0.097	0.063	0.108	0.162	0.056***	−0.194	0.393	0.203	0.490
High school graduate	−0.002	0.075	0.001	0.083	−0.017	0.043	0.101	0.328	0.113	0.388
# Children in family	−0.055	0.024**	−0.055	0.027**	−0.057	0.014***	−0.189	0.105*	−0.235	0.125*
Age of youngest child in family	0.018	0.007***	0.018	0.007**	0.019	0.004***	0.046	0.028	0.085	0.033**
Industry controls										
Retail	−0.080	0.085	−0.081	0.095	−0.077	0.048	−0.474	0.358	−0.642	0.429
Manufacturing	−0.104	0.091	−0.102	0.101	−0.111	0.051**	−0.427	0.375	−0.615	0.458
Service	−0.177	0.079**	−0.172	0.088**	−0.195	0.045***	−0.517	0.334	−0.772	0.402*
Construction	−0.098	0.117	−0.088	0.129	−0.157	0.067**	−0.115	0.477	−0.417	0.587
Duration dummy variables										
Unemployment spell began 1991	0.543	0.065***	0.563	0.072***	0.477	0.038***	2.070	0.269***	2.815	0.328***
Unemployment spell began 1992	0.562	0.078***	0.580	0.086***	0.499	0.045***	2.193	0.332***	2.892	0.393***

(*continued on next page*)

Table A.1 (*continued*)

Variables	Parameterization									
	Weibull		Exponential		Rayleigh		Logistic		Normal	
	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error	Coef.	St. error
Location controls										
Central city	−0.103	0.074	−0.095	0.082	−0.171	0.043***	0.004	0.293	−0.381	0.363
% White non-Hispanic	−0.208	0.150	−0.193	0.165	−0.286	0.086***	−0.266	0.615	−0.797	0.747
% Poor	0.104	0.336	0.081	0.371	0.267	0.192	−0.741	1.446	0.397	1.696
% High school education	0.049	0.314	0.036	0.347	0.155	0.180	−0.603	1.264	−0.190	1.542
Zip code share of total MSA jobs	−1.142	0.347***	−1.122	0.385***	−1.257	0.195***	−2.694	1.405**	−3.570	1.825*
Scale	0.901	0.025***	1.000		0.500		2.160	0.029***	4.440	0.024***
Log-likelihood	−2227.900		−2236.200		−2619.900		−2528.200		−2618.400	
AIC	4530.800		4544.400		5311.800		5131.400		5311.800	
χ^{2a}	149.200	***	134.000	***	438.800	***	118.400	***	139.600	***

[a] LR test of model significance.
* $p < 0.10$.
** $p < 0.05$.
*** $p < 0.01$.

Table A.2
Coefficients from models with alternative job accessibility controls[a]

Job accessibility controls	Coef.	St. error
Zip code share of total MSA jobs (from Model 4, Table 3)	−1.142	0.347***
Zip code share of total MSA manufacturing jobs	−0.924	0.268***
Zip code share of total MSA retail jobs	−0.936	0.339***
Zip code share of total MSA service jobs	−0.840	0.320***
Zip code total jobs	−8.2E-06	5.2E-06
Zip code jobs/worker	−0.046	0.026*
% > 25 min commute time, worker's census tract	0.109	0.213
% Work within county, worker's census tract	−0.233	0.168
% Public transit, worker's census tract	0.315	0.273

[a] All models include full set of controls shown in Model 4, Table 3, substituting variables above for "zip code share of total MSA jobs."
* $p < 0.10$.
** $p < 0.05$.
*** $p < 0.01$.

1992, but these measures were never statistically significant. As suggested by Shen [20, 21], job accessibility measures based on job changes reflect opportunities resulting from new jobs but fail to capture opportunities resulting from job turnover. The latter actually accounts for the great majority of job openings in a typical US city.

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