Comment

Teenage Employment and the Spatial Isolation of Minority and Poverty Households

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ABSTRACT

This paper tests the importance of the spatial isolation of minority and poverty households for youth employment in large metropolitan areas. We estimate a model relating youth employment probabilities to individual and family characteristics, race, and metropolitan location. We then investigate the determinants of the systematic differences in employment probabilities by race and metropolitan area. A substantial fraction of differences in youth employment can be attributed to the isolation of minorities and poor households. Minority youth residing in more segregated cities or cities in which minorities have less contact with nonpoor households have lower employment probabilities than otherwise comparable youth. Simulations suggest that these spatial effects explain a substantial fraction of the existing differences in youth employment rates by race.

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I. Introduction

Many have argued that the concentration of poor and minority households in central portions of metropolitan areas exacerbates a host of urban problems—ranging from the low quality of public services, such as education, to the high level of antisocial activity, such as violent crime. The hard evidence on the existence of concentration effects upon social outcomes is somewhat ambiguous (see Jencks and Mayers 1990 for a review; Case and Katz 1991 and Plotnick and Hoffman 1995 for recent developments), but the emergence of an urban "underclass" has generated new debate about the implications of the spatial isolation of poor and minority households upon their own well being and that of others.

Regardless of the overall effects of concentrated poverty on social outcomes, there is reason to anticipate specific impacts on the operation of urban labor markets. The well-known "spatial mismatch theory" suggests that minority workers concentrated in central cities will experience lower employment rates than will similar workers who are not spatially isolated from emerging job concentrations at suburban sites. Again, empirical evidence on the magnitude of the mismatch in jobs is not definitive (see Kain 1992 and Holzer 1991 for recent reviews), but there can be little doubt that job movement to the suburbs reduces employment opportunities for those left behind. Several recent studies have documented the relationship between the lower employment levels of black and hispanic workers and measures of travel times to jobs (Ihlanfeldt and Sjoquist 1990; Ihlanfeldt 1993).

Regardless of the importance of the "mismatch" hypothesis, the *social* isolation arising from concentrations of poverty households may itself present a barrier to employment (Wilson 1987; O'Regan and Quigley 1991). Direct observation on job search strategies indicates that a large fraction of job seekers obtain information on specific jobs from friends and relatives (Holzer 1987; Ihlanfeldt forthcoming). The importance of these informal networks in affecting access to employment suggests that some networks are far more valuable than others in obtaining employment, in other words, networks which include a larger fraction of employed members, or members with "better" jobs. Formal models of job search suggest that those in networks with low employment rates may be further disadvantaged in the labor market (Montgomery 1991; O'Regan 1993).

This paper provides an empirical test of the importance of these phenomena. We find evidence that the spatial isolation of minority and poor households contributes to differences in youth employment by race and ethnicity. Our simulations suggest that these employment effects are quite large.

II. Data and Measurement

Our empirical work is based on 1980 and 1990 Census data for non-Hispanic white (white), non-Hispanic black (black), and Hispanic youth living at home (with at least one parent) and aged 16 to 19.[1]

1. The data for 1980, a subset of the Public Use Micro Sample, cover 47 of the largest metropolitan statistical areas with 55,393 observations. The 1990 sample includes these same 47 MSAs and all other

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We focus on the employment of youth to control for the endogeniety of residential location. Several recent papers analyzing neighborhood and peer effects on behavior have highlighted the difficulty of controlling adequately for family characteristics and choice in identifying neighborhood and peer influence (Corcoran et al. 1992; Evans et al. 1992; Plotnick and Hoffman 1995). We recognize that these problems are not eliminated by focusing on youth employment. Nevertheless, in contrast to the analysis of adult workers, we can presume that the residence sites of youth are influenced less by the accessibility demands of youth and determined more by those of the family. Thus, youth primarily seek employment whose accessibility is measured from their predetermined residential locations. In addition, Census data on at-home youth include extensive data on the household in which they reside, permitting us to control for a variety of frequently omitted family characteristics.[2] This also helps to control for the endogeneity of residential location. Despite this, there may still be other omitted family characteristics that are simultaneously correlated with location and youth behavior and for which this study cannot control accurately.

In measuring racial and poverty concentration, we are specifically interested in the extent to which minorities and the poor are likely to have social contact or access to other groups. This specific dimension of segregation is best captured by a standard segregation index, the "exposure index," which measures interaction between groups.[3] It is calculated as follows:

$$(1) \quad E_{ij} = \sum_t (n_{it}/N_i)(n_{jt}/N_t).$$

E_{ij} is the exposure of the ith group to members of group j. n_{it} and n_{jt} are the number of group i and group j people in tract t, N_i is the total number of group i people in the MSA, and N_t is the total number of people in tract t. The index number, which ranges from 0 to 1, measures the probability, for the average member of group i, that a randomly picked resident of his or her census tract is a member of group j.

Social isolation of minority households decreases their contact with both nonminority (white) and nonpoor households. We presume that exposure to whites, who have higher employment rates (and perhaps greater influence in workplace decisions), is a measure of access to job contacts and, hence, to jobs.[4] We also presume that exposure to poor individuals, who are less likely to provide valuable information about jobs, indicates less access to jobs.

The measure of exposure to whites in 1980 is taken from Douglas Massey and Nancy Denton (1987); we reproduced this measure for 1990 using the same methodology. For each MSA, we calculated the exposure to whites of three

MSAs of equivalent size, 73 MSAs in total and 243,138 observations. The data for 1990 is not a public use sample. It was gathered internally at the Bureau of the Census and includes all records on youth not enrolled in school.

2. For reasons of confidentiality, the data do not include the census tract of residence.

3. See White (1986) for a comparison of various measures.

4. For the 1980 sample, the average index value of exposure to whites is: 0.870 for whites, 0.385 for blacks, and 0.668 for Hispanics. For 1990: 0.821 for whites, 0.383 for blacks, and 0.668 for Hispanics.

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groups: the exposure to whites experienced by whites; exposure to whites by blacks; and exposure to whites experienced by hispanics. The second index measures exposure to poor individuals. Using data provided by Douglas Massey and by the U.S. Census, we calculated indices of exposure to poverty for whites, blacks, and hispanics, for each MSA.[5]

III. Empirical Models

The first stage of the analysis is based on a logit model, relating youth employment probabilities to a vector of individual and family characteristics, X. The model includes race and ethnicity-specific effects which vary by MSA:

$$(2)\quad \log[P_i/(1-P_i)] = \alpha X_i + \sum_j \beta_{1j} w_i M_j + \sum_j \beta_{2j} b_i M_j + \sum_j \beta_{3j} h_i M_j$$

M_j is a set of MSA dummy variables, having a value of one if individual i resides in metropolitan area j and zero otherwise. This vector is interacted with a series of race/ethnicity dummy variables: w_i is a dummy variable with a value of one for whites and zero otherwise, b_i is a dummy variable with a value of one for blacks and zero otherwise, and h_i is a dummy variable with a value of one for hispanics and zero otherwise.

The set of parameters β_{rm} (for $r = 1, 2, 3$ races and $m = 1, 2, \ldots, 47$ or 73 metropolitan areas) represents the shift in the logit of employment probability depending upon the race of the individual and the metropolitan area in which that individual resides.

In the second stage we analyze the determinants of these metropolitan wide differences:

$$(3)\quad \beta_{rm} = \gamma Z_m + \delta E_{rm}.$$

Z_m is a vector of MSA characteristics expected to influence local labor market outcomes, and E_{rm} is the exposure index described in Equation (1). We estimate several different forms of Equation (3).

IV. Results

Table 1 presents a summary of the logit models described in Equation (2). The basic model differs from 1980 to 1990 only in the omission of the central city dummy variable, which is not available in the 1990 census. The coefficients on individual characteristics are consistent across the years. There are two exceptions. Surprisingly, in 1980, youth in female-headed households appear more likely to be employed (although the coefficient is only marginally

5. For 1980, the average index value of exposure to poverty is: 0.063 for whites, 0.194 for blacks, and 0.114 for hispanics. For 1990: 0.084 for whites, 0.216 for blacks, and 0.148 for Hispanics.

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Table 1
Logit Models of Employment Probabilities for At-Home Youth

	1980	1990
Sex (1 = female)	−0.102	−0.036
	(5.37)	(3.99)
Central city (1 = yes)	−0.100	—
	(4.33)	
Age (years)	0.274	0.335
	(21.76)	(64.53)
Education (years)	0.267	0.189
	(27.03)	(56.46)
In school (1 = yes)	−0.615	−0.504
	(24.50)	(49.89)
Female headed household (1 = yes)	0.050	−0.109
	(1.79)	(9.51)
Education of head (years)	−0.010	−0.015
	(3.10)	(8.43)
Other household income (thousands)	1.320	−1.060
	(1.89)	(11.08)
Parent working (1 = yes)	0.537	0.666
	(15.17)	(41.77)
Sample size	55,339	243,138
Chi-square	10,639.3	38,250.8
Degrees of freedom	145	208

Note: Models do not include an intercept term. *t*-ratios are in parentheses.
Model for 1980 also includes 136 dummy variables: race of the individual interacted with dummy variables for metropolitan areas (three coefficients for 47 MSAs, less five coefficients for Hispanics which could not be estimated due to small samples).
Model for 1990 also includes 200 dummy variables (three coefficients for 73 MSAs, less four coefficients for blacks and 15 coefficients for Hispanics).

significant). This result obtains only after controlling for both race and the presence of a working parent,[6] and is not found in the 1990 results.

The second difference appears in the effect of other household income (parents' and siblings') on youth employment. This result appears to reflect differences in the effect of family socioeconomic status on the behavior of youth by their school enrollment status. For 1990 the sample size is adequate to estimate the model separately for in-school and out-of-school youth. While other family income sig-

6. When the dummy variable indicating a working parent is omitted, the coefficient on female-headship is negative and significant. We include both variables in the results reported in the text, but have replicated the analysis omitting this variable (with essentially the same results throughout).

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nificantly *decreases* the likelihood of employment for in-school youth, it significantly *increases* employment probabilities for out-of-school youth.[7]

We also estimated these models with race/ethnicity dummies but no MSA dummies, and found the set of MSA coefficients to be highly significant in both years. The key finding is that, after controlling for individual characteristics, the employment probabilities of "otherwise identical" white, black, and Hispanic at-home youth vary substantially by MSA.

We now investigate the sources of these systematic differences in employment probabilities. The coefficients estimated from Equation (2) are the dependent variables, and we estimate models of the form of (3). Since the dependent variables in this analysis are regression coefficients (observed with sampling error), the models are estimated by generalized least squares.[8] Table 2 presents the results.

We expect that differences in employment probabilities for youth across metropolitan areas depend upon the aggregate economic conditions in these MSAs. We use the unemployment rate for white adults in each metropolitan area as a measure of general economic conditions. This variable has a highly significant, negative, and large coefficient in every version of these regressions we have explored.

Other aspects of the local economy differentially affect youth employment. We included a variety of measures of industry mix and found the fraction of MSA employment in the business service sector to be the best summary measure. We tested several other categories of variables in these regressions, all which proved to be insignificant.[9] Throughout, we permit intercepts to vary for the three groups,[10] to capture any systematic differences in youth employment probabilities by race and ethnicity.

Finally, after controlling for these other effects, we investigate the importance of exposure to whites, our first proxy for social access.[11] In Model I, the coefficient for the race-specific exposure index is constant across groups, and it is significantly positive in both 1980 and 1990. In Model II, we estimate separate

7. Note that previous research reports similar effects of *area* characteristics on youth regardless of school enrollment status. Freeman (1982) found similar effects of local economic conditions on youth employment for enrolled and not-enrolled youth; Ihlanfeldt and Sjoquist (1990) found that a measure of employment access had a similar effect.

8. The GLS procedure incorporates information about the estimated variance and covariances of the dependent variable (see Hanushek 1974).

9. First, we tested several MSA-level human capital characteristics, which proved to provide no additional information. Second, we attempted to control for transport access by: (i) using published Census information on commuting times and modes, (ii) using an index designed to measure local transit system access (see Linneman and Summers 1993), and (iii) using data from the Department of Transportation on public transportation systems. None of these measures adequately captures physical proximity between workers and jobs, and none were significant.

10. Specifically, we estimate $\beta_{rm} = I_1 w + I_2 b + I_3 h + \gamma Z_m + \delta E_{rm}$, where w, b, h are race/ethnicity dummies. I_1, I_2, and I_3 are the intercepts for whites, blacks, and hispanics, respectively.

11. As noted in the text, there remains the possibility of simultaneity between the measure of access and the outcome measure. Without improved data, we are limited to controlling for this by using at-home youth (whose residence choice is exogenous), extensive family background variables (thus, fewer unobservables), and information on parent employment status.

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Table 2
Inter Metropolitan Differences in Youth Employment Probabilities

	A. Exposure to Whites				B. Exposure to Poor			
	1980		1990		1980		1990	
	Model I	Model II	Model I	Model II	Model I	Model II	Model I	Model II
Unemployment rate (percent)	−0.111	−0.107	−0.128	−0.129	−0.086	−0.083	−0.103	−0.092
	(12.64)	(14.75)	(7.79)	(7.98)	(8.74)	(8.17)	(5.62)	(4.96)
Business services employment (percent)	−0.023	−0.038	0.012	0.014	−0.043	−0.040	−0.075	−0.076
	(2.34)	(4.15)	(0.62)	(0.70)	(4.66)	(4.22)	(3.69)	(3.82)
Intercept for								
Whites	−7.577	−7.105	−8.490	−8.523	−6.261	−6.162	−7.095	−7.006
	(27.05)	(22.39)	(46.38)	(35.71)	(29.18)	(27.72)	(48.99)	(46.35)
Blacks	−8.013	−7.950	−8.947	−9.014	−6.515	−6.616	−7.655	−7.866
	(34.24)	(34.92)	(57.94)	(55.26)	(27.86)	(24.61)	(47.82)	(45.45)
Hispanics	−7.570	−7.333	−8.543	−8.463	−6.234	−6.345	−7.213	−7.127
	(29.85)	(29.24)	(50.02)	(46.60)	(27.72)	(26.24)	(45.94)	(42.00)
Exposure to whites	1.064		1.105					
	(6.79)		(10.15)					
Exposure to poor					−5.489		−2.956	
					(10.56)		(8.34)	
Exposure to whites/poor by								
Whites		0.688		1.140		−7.836		−4.526
		(3.08)		(5.72)		(5.74)		(5.28)
Blacks		1.410		1.279		−5.218		−2.181
		(5.81)		(6.73)		(6.86)		(4.93)
Hispanics		0.920		0.949		−5.013		−3.783
		(4.20)		(5.69)		(6.19)		(6.66)
Sample size	136	136	200	200	136	136	200	200
R^2	0.565	0.582	0.820	0.821	0.556	0.558	0.799	0.808

Note: R^2 is from ordinary least squares regression. All coefficients are estimated by generalized least squares. (See text for details of estimation procedure.)
t ratios are in parentheses.

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coefficients for exposure to whites, by race.[12] For all three groups, in both years, exposure to whites significantly increases a youth's probability of being employed.[13]

Panel B presents analogous results using the poverty exposure index to measure social access with quite similar results. In each of the models, in both years, exposure to poverty has a negative effect upon the employment probabilities for otherwise identical at-home youth.

Because exposure to whites and exposure to poverty are highly correlated (between -0.822 and -0.878), inclusion of both indices in a single regression yields ambiguous results.[14] Exposure to whites "matters" in explaining inter urban variation in the employment propensities of minority teenagers. Exposure to the poor also "matters." Whether these are separate and distinct effects cannot be clearly determined.

The nonlinearity of the logit relationship makes it difficult to interpret the magnitude of these coefficients. The importance of these effects can be assessed more easily by simulation. We use the results described above to conduct several simulations of the impact of reduced segregation on the employment probabilities of youth. The results of a representative set of these simulations are presented in Table 3.

As a proxy for increased integration of employment groups, we simulate the effect of racial integration (or integration by poverty status) on youth employment probabilities. For each MSA, we calculate the exposure to whites (or the exposure to poverty) under complete integration and compute the implied employment probability for each individual. These probabilities are then aggregated across the MSAs.

Our simulation takes a limited resource (the "social access" provided by white youth) and redistributes it equally among all youth. Integration by race would increase the exposure of minority youth to whites, but would also decrease the exposure of white youth to other whites. In 1980, this spatial reallocation would lead to a 14–15 percentage point increase in black youth employment, a 5–7 percentage point increase in hispanic youth employment, and little or no decline in white youth employment. While this simulation reveals a substantial reallocation of employment, the aggregate employment rate changes by only a few percentage points, and actually increases.

The second simulation focuses on the segregation of poverty. The actual level of exposure to poverty is replaced by that which would be expected if poverty were evenly dispersed across census tracts within each MSA. Again, this realloca-

12. Specifically, we estimate,

$$\beta_{rm} = I_1 w + I_2 b + I_3 h + \gamma Z_m + \delta_1 w E_{1m} + \delta_2 b E_{2m} + \delta_3 h E_{em},$$

where w, b, h are race/ethnicity dummies. E_1 refers to whites, E_2 refers to blacks, and E_3 refers to Hispanics.

13. We have estimated the 1990 models using only youth who were not enrolled in school, with the same results.

14. For 1990, each measure is significant, and the OLS and GLS results are consistent with each other. For 1980, however, the OLS and GLS results differ greatly, and we cannot confidently distinguish effects of the two measures.

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Table 3
Estimated Youth Employment Rates with Spatial Integration

	Whites	Blacks	Hispanics	Average
A. 1980				
Actual employment rate (percent)	47.77%	23.08%	36.74%	42.37%
Integration by race: projected change (percentage points)	−1.44	+14.35	+4.68	+1.92
Integration by poverty status: projected change (percentage points)	−5.50	+12.58	+4.29	−1.38
B. 1990				
Actual employment rate (percent)	53.30%	29.50%	40.60%	49.00%
Integration by race: projected change (percentage points)	−3.16	+10.60	+2.27	−0.56
Integration by poverty status: projected change (percentage points)	−3.57	+4.96	+2.76	−1.73

Note: Projections are based on coefficients reported for Model II, Table 2. Due to the nonlinearity of the logit, predictions at the point of means do not equal the actual employment rates.

tion decreases minority exposure to poverty, and increases white exposure to poverty. Minority youth employment rates would increase—by 4 or 5 percentage points for hispanics, and by 12 or 13 percentage points for blacks. White employment rates would decrease by 4 to 6 percentage points, and the aggregate employment rate for youth would decline by something less than 2 percentage points. The simulations for 1990 show a similar pattern of redistribution, with barely any change in the aggregate unemployment rate.[15]

V. Implications and Conclusions

The results of this analysis provide empirical support for the employment effects of spatial segregation: The employment prospects of otherwise identical at-home youth depend, not only on the general economic conditions in the metropolitan areas in which they reside, but also on the patterns of isolation and segregation by race and by poverty status. Exposure to whites increases the

15. We have conducted these simulations for Model I, with results similar to those presented in Table 3.

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employment probabilities for youth, while residential exposure to the poor reduces employment probabilities.

Our simulations suggest that the quantitative effect of isolation on youth employment is quite large. For the simulations presented, approximately 21 to 25 percent of the existing employment gap between white and Hispanic youth is attributable to the spatial isolation of hispanics. Approximately 30 to 35 percent of the employment gap between white and black youth arises from the spatial isolation of blacks. Moreover, there is good reason to believe that these simulations underestimate the effects of concentrations on employment.[16]

Our findings are consistent with those recent studies based on neighborhoods or single metropolitan areas which have found evidence of spatial effects. For example, using geographic units approximately equivalent to census tracts, Crane (1991) found significant neighborhood composition effects on teenage pregnancy and school dropout rates. Case and Katz (1991) focused on distinct neighborhoods within one metropolitan area and found that neighborhood peers substantially influence a variety of youth behavior, including youth propensity to work. Our results are also consistent with recent work by Ihlanfeldt and Sjoquist (1990) which focused specifically on the effects of nearness to jobs upon youth employment in a single metropolitan area. Using census-tract-based measures of job proximity, they found that between 33 to 54 percent of the gap between black and white youth employment rates is explained by differential accessibility—numbers which are similar in magnitude to our results.

Given the high correlation between social and spatial access, our empirical work cannot confirm that either is a more important mechanism connecting youth to jobs.[17]

Regardless of the specific mechanism that relates youth employment outcomes to the spatial configuration of labor markets, these results document an important connection. In addition to human capital and general economic conditions, youth employment probabilities also depend on spatial isolation, and these latter factors work to the disadvantage of minority youth.

Thus, in cities with particularly isolated minority and poor populations, even modest changes in spatial isolation of these populations would dramatically improve their employment prospects.

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16. Note, for example, that the simulations assume that changing poverty and/or racial concentrations will not lead to *any* endogenous changes in youths' education attainments, or fertility behavior, or in the employment status of their parents.

17. Note, however, that while the "mismatch hypothesis" relates principally to minority households, whose residential choices are constrained by racial discrimination in the housing market, the "social network hypothesis" applies to white workers as well. Our findings are consistent with a spatial explanation that applies to all youth.

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