RACIAL AND ETHNIC GAPS IN MALE EARNINGS IN A BOOMING URBAN ECONOMY

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According to labor market queuing theory, instead of competing against one another on the basis of the wages they are willing to accept, individuals compete against one another for job opportunities (or job slots) at fixed wages. In jockeying for positions in the job queue, success for the individual worker rests on being able to signal to potential employers that he or she possesses a set of attributes (including credentials, behaviors, and attitudes) that firms find desirable. In Lester Thurow's conception of the queue, "one set of factors determines an individual's *relative* position in the labor queue; another set of factors, not mutually exclusive of the first, determines the actual distribution of job opportunities in the economy. Wages are paid based on the characteristics of the job in question, and workers are distributed across job opportunities based on their relative position in the labor queue. The most preferred workers get the best jobs" [Thurow, 1975, 76]. At the very end of the queue are those who spend little time working at all. Others find work, but because of low wages or a combination of poor pay, part-time jobs, and intermittent work, have trouble making enough income to lift themselves and their families above the poverty line. These are the "working poor."

The number of job slots available to be filled in any queue depends, for the most part, on macroeconomic factors. As aggregate demand increases, employers must go deeper into the labor queue in order to fill vacancies. As a result, only when aggregate demand is especially high do employers hire from the end of the queue — hiring those who ordinarily would be unemployed or out of the labor force altogether.

This model is at least implicitly behind research by William Julius Wilson and others who depict inner-city labor markets as being "jobless ghettos" [Wilson, 1987; 1997; Kasarda, 1990]. Blacks, and particularly black men with limited schooling, living in high-poverty neighborhoods, have been found to have extremely high unemployment rates and low labor force participation. Commentators on the urban ghetto scene often give the impression of an "underclass" so far outside the mainstream of the American economy that their attachment to the regular economy is tenuous at

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best. Much of this research is based on urban areas which were undergoing "deindustrialization" during the 1980s, such as Chicago, where Wilson did much of his field work.

A queuing model would suggest that we would not find the same tendency toward tenuous labor market attachment among minority workers with limited education in urban areas typified by strong aggregate demand. In such cities, one would expect that employers would need to go quite deep into the labor market queue in order to meet their labor needs. In labor markets where the aggregate unemployment rate is low, one would expect to find high jobless rates among inner city residents only if this labor force was so devoid of technical and social skills that employers' training costs became prohibitively high or if employers practiced racial or ethnic discrimination, whether of a pure or statistical variety.

The empirical question, then, is how do racial and ethnic minorities with limited education fare in a labor market that has had low aggregate unemployment for some time? Since 1984, the unemployment rate for the Greater Boston Consolidated Metropolitan Statistical Area (CMSA) has averaged only 4.8 percent. Except for the 1991-93 recession period, the rate has never been above 6 percent. Moreover, from 1984 until 1990, the rate never exceeded 4 percent and was as low as 2.7 percent in 1987. In such a labor market, one would expect employers to reach deep into the workforce queue to come up with the labor supply they need.

Despite an overall tight labor market, Boston, like other communities, has been undergoing a major occupational and industrial restructuring which generally favors highly skilled workers to the disadvantage of those with limited schooling. Blue-collar employment in the Boston CMSA has declined from 42 percent of all jobs to just 19 percent between 1950 and 1990 [Department of Employment and Training, 1997]. Meanwhile, professional, technical, executive, and managerial employment has nearly doubled, from 22 to 39 percent. This shift would be more problematic if it were not for the rapid expansion of low skilled health care occupations in the region including orderlies, aides, and janitors.

In this paper, we develop and test a model of labor market outcomes for black, Hispanic, and white male adults (age 21-65) with no more than a high school education living in the Greater Boston CMSA. The data for testing the model come from the *Greater Boston Social Survey* (GBSS), which includes 1,820 household interviews (with oversamples from black and Hispanic neighborhoods) carried out between the summer of 1993 and the end of 1994, when the region's unemployment rate was trending downward toward 5.0 percent from a recession high of 8.0 percent in 1992. Every effort was made to assure a representative sample of all non-institutionalized residents of the CMSA. Obviously, it is possible that some of the most disadvantaged, particularly in the inner city, were not reached. Moreover, the sample does not include those who were in prison or in other institutions at the time of the survey. The survey does not include teenagers (age 16-20) who might be expected to have the weakest labor market attachment and poorest employment outcomes. Notwithstanding these exclusions, the GBSS sample includes a population not unlike that studied by Wilson in his analysis of "jobless ghettos."

This paper has five sections. In the first section we investigate directly the labor force participation and employment rates for non-college black, Hispanic, and white men in the Greater Boston labor market. We then develop an "expected annual hours and earnings" model and combine all its components — labor force participation rates, unemployment rates, weekly hours, and hourly wages. We continue by using logit and OLS regression to study what factors are responsible for the existing variance in labor force participation rates, unemployment rates, weekly hours, and hourly wages. We then use simulation techniques to assess what role human capital, different job attributes, and discrimination play in explaining the gaps in expected annual earnings between blacks and whites and between Hispanics and whites. Finally we end with a summary of our conclusions.

LABOR FORCE PARTICIPATION, EMPLOYMENT RATES, AND HOURLY WAGES IN THE GREATER BOSTON LABOR MARKET

Based on the research of Wilson and others, one would expect the labor force participation rate (LFPR) of black men with limited education (and perhaps Hispanic men as well) to be lower than the participation rate of white men. However, in a very tight labor market this might not be true as employers move further down the labor market queue. As Table 1 demonstrates, in the Greater Boston labor market in the mid-1990s, the LFPRs are virtually identical. *Black and Hispanic men in Greater Boston were just as likely to be in the labor force as white men.* Better than 85 percent of black and Hispanic men with no more than a high-school education reported that they were in the labor force. If anything, less well-educated men of color were *slightly more likely* to be in the labor force than comparable white men. When we further restricted our sample to black men with no more than a high-school diploma who live in census blocks that are majority black, the labor force participation rate remained quite high — above 82 percent. Even after restricting the sample further to those living in census tracts where more than a quarter of the households live in poverty, nearly 70 percent of black men with limited education were in the labor force. This is one indication that the "jobless ghetto" description is not an apt one — at least not for Boston.

It is possible that each race/ethnic group could have the same proportion participating in the labor force as we have indeed found, but still have very different probabilities of actually working. In this case, the image of the "jobless ghetto" would still survive. To test this, we asked each respondent in the GBSS if they had worked anytime during the twelve months preceding the survey. Again, as Table 1 amply demonstrates, there is no substantial difference in the probability of holding a job during the year across race and ethnic groups. Even for black men living in a predominantly minority neighborhood and one where more than a quarter of the families are in poverty, the likelihood of having worked at least sometime during the year was no less than 70 percent. This still leaves nearly a third of the potential black male labor force in the "ghetto" outside the world of work — but this is a far cry from all or even most of them — and most black men in Boston, even those with little education, do not live in such heavily impoverished neighborhoods.[1]

TABLE 1
Baseline Labor Market Components

Men - Education: High School Diploma or Less (Age 21-65)
Greater Boston Labor Market: 1993-1994

	Black Men	Hispanic Men	White Men
Labor Force Participation Rate	86.2%	87.2%	82.7%
Employed in Last 12 Months	78.7%	76.8%	74.1%
Unemployment Rate	15.2%	7.1%	7.1%
Average Hours Worked per Week	34.92	43.00	50.53
Expected Annual Hours	1,327	1,811	2,020
Median Hourly Wage	$9.62	$8.14	$11.54
Expected Annual Earnings	$12,762	$14,744	$23,295
Ratios:	Black/White	Hispanic/White	Black/Hispanic
Median Hourly Wage	83.4%	70.5%	118.2%
Expected Annual Earnings	54.8%	63.3%	86.6%

Source: Analyis of Greater Boston Social Survey, University of Massachusetts Boston.

Black, and especially Hispanic, men with limited education are nonetheless still subject to an hourly earnings gap relative to white men. Referring again to Table 1, the median hourly wage for black men was only $9.62 compared to $11.54 for white men. Hence, black men earn only about 83 percent of the median of whites. Hispanic wages lag even further behind at $8.14 — only 71 percent of the white median.

On an annual basis, then, assuming all three race/ethnic groups worked full-time full-year (40 hours per week and 52 weeks per year), the annual earnings for the three groups would be as follows: Hispanic Men: $16,931; Black Men: $20,010; and White Men: $24,003. The labor market "pecking order" would clearly favor whites and most adversely affect Hispanics. Nonetheless, the full-time/full-year median annual wage would be enough for each of these groups, on average, to raise a family of four above the poverty line ($15,192). *If* they were equally able to work full-time/full-year, the median worker in each of these race-ethnic groups would not be considered, at least officially, "working poor."

All of this could be considered "good" news — at least considered against the standard of a "jobless ghetto." Unfortunately, further analysis of the GBSS data indicates that despite such high labor force participation and employment rates, non-college black men in particular are doing much more poorly than these full-time/full-year simulations would suggest.

"EXPECTED" ANNUAL HOURS AND EARNINGS

Given similar employment rates, one might conjecture that the unemployment rates would also be similar. But this does not turn out to be the case, as shown again in Table 1. While Hispanic and white men had identical unemployment rates at the time of their GBSS interviews, the black unemployment rate was about twice as high, a finding consistent with a long-standing constant in national labor force data.

One might find this somewhat mystifying given that Boston-area black men are just as likely to have worked during the twelve months prior to participating in the GBSS survey. How can the unemployment rate of black men be double the rate for Hispanic and white men while the chances of working sometime during the year are nearly identical? The answer is that black men are substantially more likely to cycle in and out of jobs. As a result, at any single point in time, they are more liable to be found unemployed despite the fact that they were employed at least sometime during the year.

Unemployment turns out to be only part of the problem for black men. As Table 1 indicates, there is a large discrepancy in usual weekly hours among the three race/ethnic groups. This reflects differences in the incidence of part-time vs. full-time work and the degree to which individuals engage in overtime. Non-college white men typically put in a very long work week — an average of 50.5 hours. Hispanics work an average of three hours over the standard 40-hour work week. Non-college black men average less than 35 hours per week.

Probing deeper into the GBSS data provides greater detail regarding differences in weekly work hours. Nearly 25 percent of black men with a high-school education or less report that they have part-time jobs averaging just 25 hours per week or less. Only 3 percent of Hispanic men and 5 percent of white men report working so few hours. At the other end of the hours spectrum, blacks are much less likely to be working overtime. Only 5 percent of the non-college black men work more than 50 hours per week on a regular basis. Fully 30 percent of white men report working this much, as do 11 percent of Hispanic men.

"Moonlighting" — working at two or more jobs in the same week — is also much more prevalent among white men. Those with a high-school degree or less report an average of 1.24 jobs each, while blacks and Hispanics both report an average of 1.04. Nearly one-fifth of white men moonlight; less than one-twentieth of minority men do.[2] An even stronger disparity occurs when it comes to self-employment. White men are much more likely to be self-employed —often working for themselves in addition to holding down a regular job elsewhere. Nearly two out of five (38 percent) white men report self-employment income, and that often means they are working extra hours. Only 4 percent of black men and 3 percent of Hispanic men work for themselves.

We can now take all of the labor market components and combine them to yield a measure of "expected" annual hours and "expected" annual earnings for each group. These are calculated according to the following formula:

$$Y = \{\text{pr}\ (LFP) \times [1 - \text{pr}\ (UR)] \times H/Wk \times 52\ \} \times W \quad (1)$$

where: Y = expected annualized earnings
$\{\text{pr}\ (LFP\ ...52\}$ = expected annualized hours,

pr (*LFP*)	=	probability of participating in the labor force,
[1–pr(*UR*)]	=	(1 – probability of being unemployed),
H/Wk	=	mean hours worked per week,
W	=	median hourly wage.

Group annual earnings are therefore a function of the group probabilities of being in the labor force and being unemployed as well as mean weekly hours of work and measured median hourly wages.[3]

The expected annual hours for the three race/ethnic groups are shown in Table 1. Black men can expect to put in, on average, just a little over 1,325 hours per year compared with over 1,800 hours for Hispanic men and more than 2,000 hours for whites. Multiplying these values by the median hourly wage yields the expected annualized earnings found in Table 1. With an expected value of $12,762 in annual earnings, the black men in our sample fell over $2,400 — or 16 percent — shy of the official poverty line for a family of four in 1994 ($15,192). In contrast, Hispanic men were less than $500 below the poverty line, while white men had expected annual earnings more than 50 percent above.

LOGIT AND OLS ESTIMATES OF LABOR MARKET OUTCOMES

To probe deeper into the determinants of labor market outcomes for men with no college experience, we developed logit equations for labor force participation and unemployment rates, and ordinary least squares (OLS) regressions for weekly hours and hourly wages. Combining the results from these equations via the earnings formula then allows us to ascertain the role of human capital, race, and residential location on "expected" annual income. The specific methodological protocol for the model is described in Appendix A.

Table 2 provides a complete list of the final variables and their definitions used in this model. Table 3 provides the mean (or median) values for the dependent and independent variables in the final equations. Table 4 presents the entire set of final fitted equations. The sample for these equations included all men in the GBSS who were between the ages of 21 and 65 and who had formal education of 12 years or less. Sample sizes for each equation differ based on data availability. While we expected such variables as *AGE*, *AGESQ*, *TOTJEXP* (job experience), and *FORBORN* (foreign born) to have an impact on labor force participation and unemployment, their coefficients did not prove statistically significant at the .10 level.[4] Therefore, in the final simulation equations, these variables were dropped.

One variable requires additional justification: *STRAT* (the dummy variable signifying whether an individual lives in a census block where a majority of residents are black or Hispanic). While generally identified here as reflecting the impact of neighborhood composition on labor market outcomes, this variable could be a proxy for any form of "unmeasured" human capital that is shared by a large number of residents in minority neighborhoods and not prevalent in "majority" neighborhoods. Hence, we must exercise some caution in interpreting its meaning in this model.

TABLE 2
Variable List for Labor Market Simulation Model

Variable Name	Variable Description
LFP	In labor force (Dummy Variable)
UR	Unemployed (Dummy Variable)
HRSWEEK	Usual hours worked per week
RLnHWAGE	Log of real hourly wage — including employment and self-employment earnings (adjusted for time elapsed since last job)[a]
BLACK	Non-Hispanic black (Dummy variable)
HISPANIC	Hispanic (Dummy variable)
STRAT	Live in black or Hispanic majority census block (Dummy variable)
HSDEGREE	Completed 12 years of schooling (Dummy variable)
ASSESSUQ	Interviewer's assessment of ability of respondent to understand questions on survey (1=excellent to 5=poor)
HEALTHDY	Health condition limits type or hours of work (Dummy Variable)
AGE	Age
AGESQ	Age squared
TOTJEXP	Years of experience in current or last occupation
ARMFORCE	Served in either active duty or military reserves (Dummy Variable)
FORBORN	Mother living outside U.S. when born
MARRIED	Ever married (Dummy Variable)
NUMJOB	Number of jobs held at time of survey
SELFEMP	Self-employed on current or last main job (Dummy Variable)
SERVICE	Current or last main job in service occupation (Dummy Variable)
SALES	Current or last main job in sales occupation (Dummy Variable)
UNION1	Member of a craft, industrial, or trade union (Dummy Variable)
COMPJOB	Work with computer daily or weekly on job (Dummy Variable)
EMPREC	At work within past 12 months (Dummy Variable)

a. Inflation adjustment based on .2228 percent monthly increase in nominal average wages for production and nonsupervisory workers (1988-1995). Calculated from Council of Economic Advisers, *Economic Report of the President 1996*, Table B-43, 330.

Moreover, it is likely that a simultaneous relationship exists between labor market outcomes and STRAT. Those living in minority neighborhoods may see their labor market success constrained by reason of where they live. Alternatively, limited labor market success may result in low incomes which, in turn, limit the types of neighborhoods these individuals can afford. As Ihlanfeldt points out, one can deal with this problem either by estimating a system of equations which treats both employment and job access as endogenous variables or by restricting the sample to individuals for whom residential location is truly exogenous (e.g. teenagers living with their parents) [Ihlanfelt, 1992; also see Waddell, 1992; Waddell, 1993].

In our case, developing a full system of simultaneous equations, given the simulation methodology employed here, would have generated a number of intractable problems. Since our sample excludes teenagers, we could not follow the second method, either. The question, then, was whether to include *STRAT* at all. We chose to include this variable for two reasons. First, excluding this variable would have subjected the

TABLE 3
Mean/Median Values for Labor Market Simulation Model MEN
(High School or Less; Age 21-65)

Variable Name	Black	Hispanic	White
LFP	.862	.872	.827
UR	.152	.071	.071
HRSWEEK	34.9	43.0	*50.5*
RHWAGE (Median)	$9.62	$8.14	$11.54
STRAT	.782	.365	.006
HSDEGREE	.579	.282	.921
ASSESSUQ	2.475	2.902	1.589
HEALTHDY	.211	.238	.184
AGE	39	30	39
TOTJEXP (Median)	6	6	8
ARMFORCE	.162	.056	.304
FOREIGN BORN	.390	.656	.020
MARRIED	.540	.885	.734
NUMJOB	1.039	1.037	1.239
SELF EMPLOYMENT	.041	.032	.381
SERVICE OCCUPATION	.114	.215	.041
SALES OCCUPATION	.054	.010	.204
UNION MEMBER	.277	.310	.122
COMPUTER USE	.103	.179	.321
RECENTLY EMPLOYED	.787	.768	.741

analysis to specification error as a result of a missing variable. By excluding *STRAT* from the analysis, it is likely that we would have obtained biased coefficients on such variables as race and ethnicity. Moreover, in a forthcoming study, we have demonstrated that rents in Boston area minority neighborhoods are not significantly different from rents in white neighborhoods after controlling for the education level of the renters [Bluestone & Stevenson, 1999]. Hence, affordability does not appear to be a major factor constraining the decision to locate in a majority or a minority neighborhood. This suggests that while the relationship between *STRAT* and labor market outcomes may still be simultaneous, the causation appears to run primarily from residential location to labor success, not vice-versa. Minority workers are constrained as to where they live as a result of factors other than income. As is, *STRAT* was not significant in three of the four equations, suggesting that it plays only a modest role in labor market outcomes for this population.

Labor Force Participation — A logit equation was used to examine the factors that are associated with whether an individual was either currently working or at least actively seeking employment at the time of the GBSS survey.[5] Recall that more

TABLE 4
Logit and Regression Equation Results: Men
High School Degree or Less Schooling

	LFP (logit)	UR (logit)	HRSWEEK (OLS)	RLNHWAGE (OLS)
Common Variables				
BLACK	1.5683	-.0127	-4.8153	-.2101
	(1.79)	(0.01)	(2.31)	(2.13)
HISPANIC	1.2878	-.4627	5.7463	.1834
	(1.21)	(0.45)	(2.00)	(1.59)
STRAT	-.6783	1.1985	.0151	-.1723
	(0.97)	(1.329)	(0.007)	(2.77)
Human Capital Variables				
HSDEGREE	.5257		2.0639	-.0007
	(0.80)		(1.27)	(0.01)
ASSESSUQ				-.1085
				(4.07)
HEALTHDY	-3.0707		-4.0835	
	(2.36)		(2.36)	
AGE			1.9332	.0514
			(4.22)	(3.29)
AGESQ			-.0221	-.0006
			(3.95)	(2.94)
TOTJEXP				.0083
				(1.68)
TOTJEXP*BLACK				.0197
				(3.19)
TOTJEXP*HISPANIC				.0062
				(0.73)
ARMFORCE				.0019
				(0.15)
ARMFORCE*BLACK				.1736
				(1.46)
Nativity/Family Status Variables				
FOREIGN BORN			-.7462	.2978
			(0.36)	(3.21)
FOREIGN BORN*BLACK				-.0634
				(0.54)
FOREIGN BORN*HISPANIC			-5.9200	-.5224
			(1.71)	(3.91)
MARRIED	.9801			
	(1.52)			
Job Characteristic Variables				
NUMJOB			10.4207	
			(4.27)	
SELF EMPLOYMENT			6.3738	
			(2.02)	
SERVICE OCCUPATION			-4.6126	
			(2.55)	
SALES OCCUPATION			10.2931	
			(2.35)	

TABLE 4 (Cont.)
Logit and Regression Equation Results: Men
High-School Degree or Less Schooling

	LFP (logit)	UR (logit)	HRSWEEK (OLS)	RLNHWAGE (OLS)
UNION MEMBER				.2122 (2.81)
COMPUTER USE				.1167 (2.01)
Control Variable				
RECENTLY EMPLOYED				.1929 (2.07)
Constant	1.5180	-2.6874	-7.7026	1.1537
N	318	252	279	245
F	1.36	1.27	4.68	14.35
Prob > F	.234	.287		
R-Squared			.256	.354
	(weighted)	(weighted)	(unweighted)	(unweighted)

Source: *Greater Boston Social Survey.* t-statistics are in parentheses.

than five out of six men in our study reported themselves to be in the labor force with little variation by race or ethnicity. Hence, there is not much actual variation to explain. Only three variables of the six included in the equation proved statistically significant at the .10 level or better: *BLACK*, *HEALTHDY*, and *MARRIED*.

Of these variables, *HEALTHDY* and *MARRIED* have the expected signs. Those reporting a health condition which limits the type of work they can do or their hours of work were much less likely to be in the labor force. Based on a simulation of the labor force participation equation, black, Hispanic, and white men with no health limitations averaged labor force participation rates (LFPRs) of 90-92 percent. Those with health limitations have estimated LFPRs of 40, 44, and 28 percent.

The impact of marital status is not anywhere as large, but is still significant. Using the equation to simulate LFPR's, we find a 10-16 percentage point difference between single and married men.[6] Being married is positively related to being in the labor force, but we cannot ascertain from our data whether it means that married men are more motivated to work or that working men are more likely to be "marriageable."

The positive coefficient on *BLACK* was not originally expected, given the literature regarding labor market discrimination and "jobless ghettos." But it is compatible with and reflective of the surprising finding, reported earlier, that the point estimates for the labor force participation rates for both black and Hispanic men with limited education are, if anything, slightly *higher* than the estimate for white men.

Moreover, that the coefficient on *STRAT* — the variable that reflects whether an individual lives in a "majority-minority" neighborhood — was negative as expected, but *not* statistically significant, suggests that after controlling for an individual's race, living in a racially segregated neighborhood does not appear to contribute very much to reduced labor force participation in Greater Boston. That is, race and ethnicity may play a major role in producing disadvantage in the labor market for people of color, but it is not manifested through labor force participation.

We tried other variables in the model including foreign born, age, high-school completion, and veteran status, but none of these proved to be anywhere near statistically significant.[7] This was also true of variables created to proxy for transportation options which might be related to an individual's ability to commute to a job. Overall the logit equation itself has a low F-statistic, suggesting that the combination of variables used to explain labor force participation do a poor job of revealing the distinguishing characteristics of those who are in and out of the labor force. With the exception of race and marital and health status, labor force non-participation seems to be generally idiosyncratic for the less-educated men in the GBSS sample. More than 95 percent of healthy black men with limited schooling are in the labor force, as are 99 percent of the Hispanics and 93 percent of the white men.

Unemployment — Our attempt at explaining unemployment was somewhat problematic. An initial unemployment equation, conditioned on being in the labor force, and based on unweighted data proved highly unstable in simulation experiments. Moreover, it produced a rank order of jobless rates by race/ethnic group quite different from that found in Table 1.[8] These results forced us to abandon this approach and turn to the weighted data as we had in the case of the labor force participation equation. In the end, the weighted data provided for a suitably stable model which accorded with the relative rankings of unemployment rates.

Even then, none of the variables we used to ascertain the probability of being unemployed were statistically significant. The F-statistic for the equation was also quite low. The closest we came to a statistically significant variable was *STRAT* — suggesting that those who reside in minority neighborhoods may be disadvantaged when it comes to finding and retaining a job. Indeed, if we disregard the low t-statistic on this variable and insert its coefficient at face value in the unemployment rate module of the annual earnings simulation (to be discussed in the next section of this chapter), we find that moving black men from a minority neighborhood to a "white" neighborhood reduces their estimated unemployment rate from 15.2 percent to 6.7 percent — about the same as the white and Hispanic jobless rates. The same simulation reduces the Hispanic unemployment rate to just 4.4 percent from 7.1 percent. The *spatial dimension* of the labor market therefore seems to be most pronounced when it comes to unemployment, not labor force participation — or hourly wage rates as we shall demonstrate.

What was quite obvious from extensive experimentation with the unemployment equation is that many conventional variables do not seem to provide an explanation for differences in the probability of being unemployed. We expected that access to a car would help explain differences in unemployment, but this variable proved insignificant as well. Similarly, we expected that those who searched for work using a network comprised of family and friends would face lower unemployment. However, none of a series of "network" or social capital variables proved statistically significant either. In a much larger sample, these might have proven significant.

Weekly Hours — There was much greater success in explaining differences in weekly working time. While labor force participation and unemployment in the GBSS refer to current status, data on usual weekly hours and on hourly wages refer to current job *or* previous job within the past five years. This substantially eliminates the problem of selection bias since we have hours and wage data for all those with any work experience over the past five years even if they are currently not participating in the labor force or are currently unemployed. Hence, the equations for hours and wages are not conditioned on labor force status, obviating the need for implementing a procedure to correct for selection bias.[9]

A large number of GBSS variables enter the model with the expected signs and statistically significant coefficients. Altogether, our ordinary least squares (OLS) regression explains about a quarter of the variation ($R^2 = .256$). Other things equal, race and ethnicity matter. Black men work an average of five hours *less* per week than whites; Hispanics work nearly six hours *more*. After controlling for race and ethnicity, we found no evidence of a "ghetto" effect on weekly hours. STRAT has the expected sign, but its coefficient was close to zero and insignificant.

Those with a health limitation not only are less likely to participate in the labor force, but when they are working they tend to work four hours less per week than others. Working time increases with age up to age 44 and then slowly declines. Foreign born men typically work about the same time as native born.[10]

But, the most important set of factors determining working time has to do with the number and types of jobs that men with limited education obtain. As expected those who work multiple jobs work longer hours — about 10 hours more per week. This suggests that a not uncommon experience, at least for white men, is to work a regular job of 40 hours plus a part-time job averaging about 10 hours. The combination gives the mean 50.5 hour work week we found earlier for white men. Self-employment also contributes to longer work weeks. Those who are self-employed on a regular basis work 6 hours more per week than those who are not. Part of this hours differential may reflect moonlighting as well, with the self-employed working part of the week for someone else and part of the week for themselves.

Other things equal, working in service occupations reduces reported working time by nearly 5 hours per week. In contrast and somewhat unexpectedly, working in sales occupations does the opposite. It increases weekly hours by a whopping 10 hours. For those who think of K-Mart or a similar retail outlet as the locus for most sales jobs, this result might seem strange indeed. But sales occupations are found in a broad range of industries, not just retail trade. In financial services, insurance, banking, wholesale trade, and manufacturing, sales representatives make up a large share of total employment. These sales reps often work very long hours out of offices or via telemarketing. The coefficient on *SALES OCCUPATION* is therefore not counter-intuitive once the distinction between "occupation" and "industry" is made.

The one variable that seems to play a minor role in working time is the high-school degree. Those who have dropped out before completing high school suffer at most a two hour deficit in hours worked. Altogether, then, we find that race, ethnicity, nativity, health status, age, moonlighting, self-employment, and occupation all affect working time — and therefore, in the final analysis, they all affect annual earnings.

How much of this important difference in weekly hours is voluntary? Do black men choose to work fewer hours or is this an involuntary outcome? While this analysis cannot provide a definitive answer to these questions, there is strong reason to believe that racial differences in working time are not primarily a matter of individual choice. Much of this difference can be explained by the racial pattern of the occupational distribution in Greater Boston. Black men are only one-fourth as likely as white men to be in sales occupations — positions that generally offer significantly longer hours. On the other hand, black men are more than twice as likely to be in service occupations that generally offer shorter work weeks. To the extent that occupational choice is limited by various forms of racial discrimination, differences in working time is the result. Moreover, it is interesting to note that since 1983, black workers nationwide have increased their average weekly work time more than either white or Hispanic workers [Bluestone and Rose, 1998]. This suggests that motivation is not the key factor in the large hours gap we find in Greater Boston.

Hourly Wages — The final component in our labor market outcomes model is the hourly wage. The OLS regression for the natural log of real hourly earnings includes a large number of statistically significant variables related to human capital, family background, and job characteristics — as well as race and residence. Overall, the regression explains 35 percent of the total variance in wages.

Race is a highly significant variable, but it interacts in complex ways with occupational experience, veteran status, and nativity.[11] An evaluation of the regression results suggests that a native born black who has only one year of occupational experience and has not served in the military will earn about 17 percent less than a comparable white. However, as his occupational experience increases, the racial wage gap declines. With six years of experience the gap is down to 9 percent. By ten years of experience, it disappears. An alternative to occupational experience is to have served in the armed forces. A black veteran who has just begun his career and has only one year of occupational experience will, according to the regression, earn about the same hourly wage as a comparably inexperienced white, regardless of the latter's veteran status.

Significant black disadvantage therefore exists in the Greater Boston region when it comes to hourly wages, but it declines with job experience. Additional years of experience in the same occupation permit black men in particular to move up the seniority wage ladder. Black men apparently enter occupations at the very bottom rung, but those who build up seniority benefit disproportionately. Such a wage pattern is consistent with a model of statistical discrimination. Hired in at low pay, black men who prove themselves to their employers move up the wage ladder more swiftly than either white or Hispanic men. Still, the problem is getting hired in the first place and then continuing to be employed long enough in a given occupation to enjoy such benefits.

That statistical discrimination plays a large role in the labor market is also confirmed by the impact of veteran status. The variable *ARMFORCE* might be a proxy for either technical skills or social skills. But, since the variable is only important for black men, the "signaling" role of veteran status is probably more pertinent. To the extent that black men need *better* credentials to compete for the same job opportunities as whites, veteran status is a highly valued credential.

In contrast to blacks, Spanish ethnicity *per se* does not appear to confer a disadvantage in terms of hourly earnings, even though Hispanic men have the lowest hourly wage of the three race/ethnic groups in our analysis. This suggests that other factors are responsible for the low hourly earnings of Hispanic men, not ethnicity *per se*.

The variable *STRAT* enters the wage equation as statistically significant. According to its coefficient, employed men who live in minority dominated neighborhoods suffer a 16 percent wage deficit. We generated terms for *BLACK* and *STRAT* and for *HISPANIC* and *STRAT* to check for interaction effects but the coefficients on these terms proved both small and statistically insignificant. Therefore, it appears that race and ethnicity on the one hand and the neighborhood variable *STRAT* on the other have independent effects on hourly wages — although the race effect is conditioned on job experience and veteran status.

A large number of human capital and demographic factors play a role in wage determination for non-college men in the Greater Boston metropolitan region. While completing high school does not appear to have the expected "diploma" effect — its coefficient is small and insignificant — age and years of occupational experience are both significant. The normal age-earnings profile suggests that, on average, men reach their peak earning capacity at around age 43. This helps to explain why the Hispanics in our survey have lower wages than others. Their median age is only 30 compared with 39 for both blacks and whites. If the age distribution of Latino men matched that of whites, a good portion of their hourly wage deficit would disappear. An estimate of how much will be provided when we turn to the simulation of the entire labor market model.

The GBSS contains a number of questions regarding individual abilities and achievement. Among these are the results of a brief word recognition test given each respondent (in English or Spanish, depending on what language was used in the interview) and interviewer assessments of the respondent's ability to understand English, to speak English clearly, and to answer the survey questions. Of these, one factor proved highly significant in the wage equation —*ASSESSUQ*, the interviewer's assessment of the ability of the respondent to understand the survey questionnaire.[12] While this measure is clearly subjective, the high *t*-statistic and the reasonable coefficient provide some confidence in this measure.[13] Given the attempt at race and ethnic matching between interviewer and respondent, there is reason to believe that the amount of systematic bias in these assessments based on the race or ethnicity of the interviewer is small.[14] Nevertheless, the mean scores vary significantly by race and ethnicity: 1.59 for white men; 2.48 for black men; 2.90 for Hispanics.[15] Hence, while the high-school degree *per se* has no apparent role in wage determination for *this* group of men, differences in "ability" — as indicated by these assessments — seem to matter a good deal.[16]

That *HSDEGREE* was not significant in any of the component equations may seem quite surprising. What may explain this result is that in a regional economy with such a tight labor market, employers cannot afford the luxury of limiting their search to workers who are high-school graduates. Indeed, a recent news story describes Boston area employers' recruitment efforts inside homeless shelters in inner city neighborhoods [Stein, 1998]. This is not to say that schooling never counts in such a labor market. We have run other GBSS analyses in which the data set was not truncated on years of education. In this case, schooling turns out to be a critically important factor in explaining differences in annual earnings [Bluestone, Massagli, and Stevenson, 1996].

Beyond the impact of "human capital," the wage equation provides an interesting story about nativity. Evaluating the coefficient on *FORBORN* along with its two race/ethnic interaction terms suggests that foreign born white workers — with limited education — are doing substantially better than native born whites. *Ceteris paribus*, they earn about one-third more per hour (34 percent) than white men born in the United States. It may be true that those who come to the United States do so specifically for economic reasons — and working at a good job is one of them. Foreign born blacks do better than native born, as well. Immigrants, primarily from the Caribbean and parts of Africa, are doing better than those who were born in the United States. The fact that nearly two out of five (39 percent) black men in the region with limited schooling are foreign born makes this finding particularly salient. In the ordering of the labor market, whites dominate blacks and Hispanics, but foreign born blacks dominate those who were born and raised here. The mean hourly wage of non-college foreign born blacks is $10.52 versus $9.56 for native born. According to the regression, the opposite pattern is true among Hispanics. Those who have recently come to the United States from abroad (or from Puerto Rico) would earn, if everything else were held constant, 25 percent less than Latino men who have lived here all their lives. This may reflect differences in language ability that we were not able to test adequately.

As for the impact of job characteristics on the wage, two variables were found to be statistically significant (*UNION MEMBER* and *COMPJOB*). Those who are members of unions in the Greater Boston region receive a 24 percent wage premium. This is consistent with national data which demonstrate that the union/nonunion gap is generally of this magnitude [Mishel, Bernstein, and Schmitt, 1999].[17] Black and Hispanic men particularly benefit from unions in Boston because approximately 30 percent of those with high-school diplomas or less are union members. Whites more often work in industries and occupations where unions are much less prevalent. The white unionization rate in our weighted sample is only 12 percent. According to our regression, if it were not for relatively high union density among black and Hispanic men, the racial and ethnic wage gaps would be significantly larger.

Those who work regularly with computers also receive a wage premium, on the order of 12 percent.[18] Unlike unionization, however, this wage premium favors white men. About a third (32 percent) of the non-college white men in our study reported working with computers on their jobs on a daily or at least weekly basis. Among Hispanics, only about half this proportion used computers, and among black men only one in ten. Clearly, access to jobs requiring computers is an important determinant of wages. Black men are least likely to have such an opportunity for labor market success.

The final variable in the hourly wage equation is a control variable, *EMPREC*, which indicates whether the sample individual had actually worked at any time during the previous twelve months. Because hourly wages in the GBSS are available not only for currently working men, but for those who have worked any time in the past five years, it is necessary to check for differences in wages that might be associated with current work status. The positive and significant coefficient on the control variable indicates that those who have worked at least sometime in the past year earn, *ceteris paribus*, 21 percent more than those who reported no work in the past year. This differential could be due to any number of factors including differences in human capital that are not otherwise captured in the regression equation.

SIMULATING LABOR MARKET OUTCOMES

The final step in our inquiry is to bring all of this information together into a single simulation model so that we can evaluate how each of these factors affect "expected" annual earnings. To accomplish this, we entered the four labor market component equations into a spreadsheet and then linked the equations by means of the labor market outcome formula reported earlier as Equation (1). By varying values of the "independent" variables in the model, we can simulate the impact on hourly wages and annual earnings. The complete baseline model is depicted in Table 5.[19] The best way to summarize the simulation results is by way of a series of charts developed from the spreadsheet models.

Human Capital Factors

How much of the race/ethnic earnings gap could be closed if differences in education, age, health status, occupational experience, and veteran status were eliminated? Figure 1 provides the answer to this question. Inserting the white male human capital values into the simulation model reduces the hourly wage gap for both blacks and Hispanics — with slightly more improvement for Hispanic men. Since black men have the same median age as white men, black hourly wages do not improve due to substituting the white median. But for Hispanics, the earnings ratio increases to .80 from .71 — and it is now within just three percentage points of the black/white ratio. Because the coefficient on *HSDEGREE* is small (and insignificant), the impact of awarding blacks and Hispanics diplomas in the same proportion as whites has little impact on closing the wage gap. The same thing is true for health status: while we found that adverse health conditions limit working time, they do not appear to have

TABLE 5
Factors Contributing to Male Annual Earnings Differences
High School Degree or Less

Men: Baseline Simulation

LABOR FORCE PARTICIPATION (Used Weighted Logit Equation)

Logistic Regression	b	Simulations: Black	Hispanic	White	All
BLACK	1.5683	1	0	0	0.071
HISPANIC	1.2878	0	1	0	0.097
HEALTHDY	-3.0707	0.211	0.238	0.184	0.187
MARRIED	0.9801	0.540	0.885	0.734	0.747
STRAT	-0.6783	0.782	0.365	0.006	0.097
HSDEGREE	0.5257	0.579	0.282	0.921	0.836
Adjusted Constant	0.9050				
Original Constant	1.5179				

Simulation: Predicted Value:	Prob.	Mean	Adj.Prob	BM/WM	HM/WM
Black	0.894	0.862	0.862	104.1%	105.4%
Hispanic	0.903	0.872	0.872		
White	0.823	0.827	0.827		
All	0.842	0.842			

UNEMPLOYMENT

	b	Simulations: Black	Hispanic	White	All
BLACK	-0.0127	1	0	0	0.071
HISPANIC	-0.4627	0	1	0	0.097
STRAT	1.1985	0.782	0.365	0.006	0.097
Adjusted Constant	-2.6300				
Original Constant	-2.6874				

Simulation: Predicted Values	Prob.	Mean	Adj.Prob	BM/WM	HM/WM
Black	0.154	0.152	0.152	2.15	1.00
Hispanic	0.066	0.071	0.071		
White	0.068	0.071	0.071		
All	0.072	0.072			

HOURS/WEEK

	b	Simulations: Black	Hispanic	White	All
BLACK	-4.815	1	0	0	0.071
HISPANIC	5.746	0	1	0	0.097
STRAT	0.01509	0.782	0.365	0.006	0.097
HEALTHDY	-4.083	0.211	0.238	0.184	0.187
NUMJOBS	10.420	1.039	1.037	1.239	1.19
AGE	1.9332	39	30	39	38

TABLE 5 (Cont.)
Factors Contributing to Male Annual Earnings Differences
High School Degree or Less

Men: Baseline Simulation

HOURS/WEEK Simulations:

	b	Black	Hispanic	White	All
AGESQ	-0.02208	1521	900	1521	1444
SERVICE OCC	-4.6125	0.114	0.215	0.041	0.064
SALES OCC	10.2931	0.054	0.01	0.204	0.163
SELFEMP	6.3738	0.041	0.032	0.381	0.297
HSDEGREE	2.0639	0.579	0.282	0.921	0.836
FORBORN	-0.7461	0.39	0.656	0.02	0.167
HISP FB	-5.9199	0	0.656	0	0.064

Adjusted Constant -7.7026
Original Constant -4.8428

Simulation: Predicted Values	Predicted	Mean	Adj.Pred	BM/WM	HM/WM
Black	40.46	34.90	34.90	0.691	0.852
Hispanic	41.53	43.00	43.00		
White	52.49	50.50	50.50		
All	50.22	50.22			

LNHOURLY WAGE Simulations:

	b	Black	Hispanic	White	All
BLACK	-0.2100	1	0	0	0.071
HISPANIC	0.1833	0	1	0	0.097
STRAT	-0.1723	0.782	0.365	0.006	0.097
EMPREC	0.1928	0.787	0.768	0.741	0.76
AGE	0.0514	39	30	39	38
AGESQ	-0.0005	1521	900	1521	1444
COMPJOB	0.1167	0.103	0.179	0.321	0.323
ARMFORCE	0.0011	0.162	0.056	0.304	0.252
BLARMF	0.1736	0.162	0	0	0.012
TOTJEXPA	0.0082	6	6	8	8
UNION1	0.2121	0.277	0.31	0.122	0.144
HSDEGREE	-0.0007	0.579	0.282	0.921	0.836
BLEXP	0.0197	6	0	0	0.426
HISEXP	0.0062	0	6	0	0.582
FORBORN	0.2978	0.39	0.656	0.02	0.167
BLACKFB	-0.0634	0.39	0	0	0.028
HISPFB	-0.5224	0	0.656	0	0.064
ASSESSUQ	-0.1084	2.475	2.902	1.589	1.8365

Adjusted Constant 1.2206
Original Constant 1.1536

TABLE 5 (Cont.)
Factors Contributing to Male Annual Earnings Differences
High School Degree or Less

Men: Baseline Simulation

LNHOURLY WAGE
Simulation:
Predicted Values

	LNHRWAGE	Median (LNHRWAGE)	Adj. (LNHRWAGE)	**BM/WM**	**HM/WM**
Black	2.2536	2.2630	2.2630	**92.5%**	**85.7%**
Hispanic	2.2291	2.0970	2.0970		
White	2.4618	2.4460	2.4460		
All	2.4460	2.4460			

	exp(LnHW)	Median	Adj.	**exp(LnHW) BM/WM**	**HM/WM**
Black	9.52	9.62	9.62	83.4%	70.6%
Hispanic	9.29	8.14	8.14		
White	11.73	11.54	11.54		

ANNUAL HOURS estimated from equations

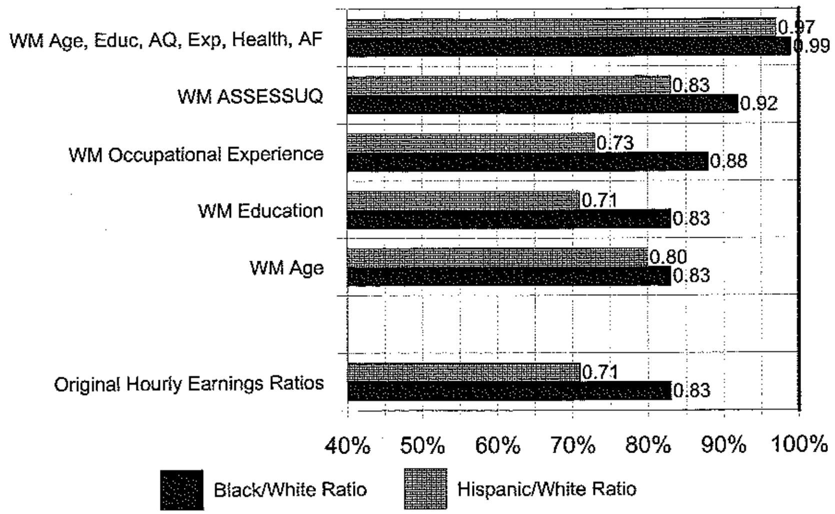
	Baseline	Simulated	**BM/WM**	**HM/WM**
Black	1,327	1,326	**65.7%**	**89.7%**
Hispanic	1,811	1,812		
White	2,018	2,019		

an independent impact on wages. Substituting the white male median for occupational experience does little for Hispanics, but raises the black ratio to .88. The combination of white male education and *ASSESSUQ* raises the ratios to .83 and .92 for Hispanics and blacks, respectively.

Finally, when we substitute white male age, education, *ASSESSUQ*, occupational experience, and veteran status, both the black/white and Hispanic/white ratios close nearly to unity — to .99 and .97 respectively. As a result, after simulating equivalent human capital investments in all three groups of non-college men, we find virtually no difference in hourly wages among white, black, and Hispanic workers. Any wage effect that might be deemed due to "discrimination" (as evidenced by the negative and statistically significant coefficient on the variable *BLACK*) is offset by "advantages" black men have in union membership and the apparent strong positive signals of occupational experience and veteran status. This particular finding of racial convergence in hourly wages after controlling for human capital is consistent with a growing literature on the black/white *hourly* wage gap [Ferguson, 1993; O'Neill, 1990].

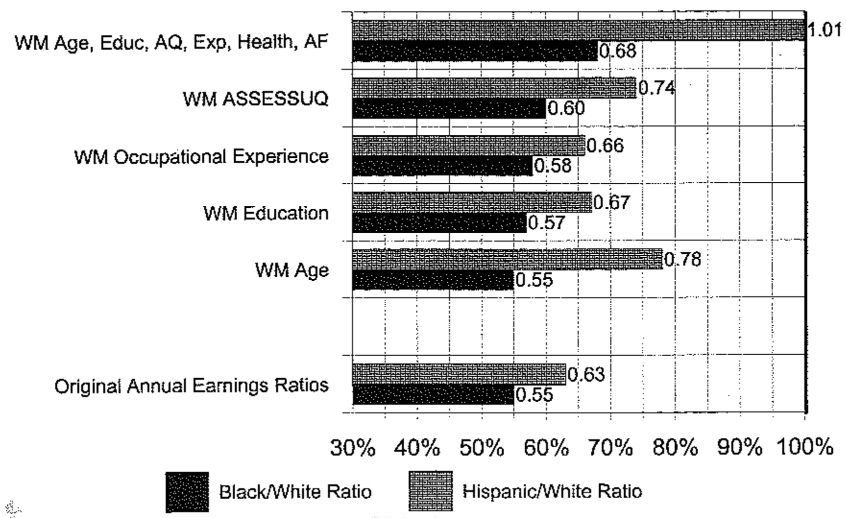
But simulating the entire model for expected *annual* earnings gives very different results as Figure 2 amply demonstrates. Recall that the original annual race/ethnic earnings ratios are .55 and .63 for blacks and Hispanics respectively. Assigning the white median age to black men does nothing to close the earnings gap because both racial groups have the same median age. But the generally younger Hispanic

FIGURE 1
Men: Hourly Earnings Ratios
Impact of Human Capital Factors



Population of GBSS men with High School Degree or Less Schooling

FIGURE 2
Men: Annual Earnings Ratios
Impact of Human Capital Factors



Population of GBSS men with High School Degree or Less Schooling

men gain substantially in this simulation, closing the Hispanic/white gap from .63 to .78. This occurs because as young workers enter middle age *both* weekly hours and the hourly wage increases. As we artificially age Greater Boston's Hispanic population, two-fifths (40 percent) of the Hispanic/white male annual earnings gap of 37 percentage points disappears.

The completion of high school improves the black/white ratio slightly (from .55 to .57), and helps the generally less-educated Hispanic men somewhat more, closing the earnings gap by 4 percentage points. Equalizing occupational experience closes the gap a little bit as well. The combination of white male values for education and *ASSESSUQ* closes the gap still more, to .63 for black males and to .79 for Hispanics. Yet only when we simulate the model giving black and Hispanic men the entire set of white male human capital traits does the overall impact of human capital variables become dramatic. For black men the earnings gap closes by 13 percentage points (from .55 to .68) leaving black men still one-third (32 percent) behind the annual earnings of white men. On the other hand, the combined human capital injection raises the Hispanic/white male ratio to 1.01. *For Hispanics, closing the human capital gap eliminates the original annual earnings deficit altogether. Raising Hispanic men to the age, education, occupational experience, ASSESSUQ, and health and veteran status of white men leaves them at parity with white men in earnings.* This is not true for black men.

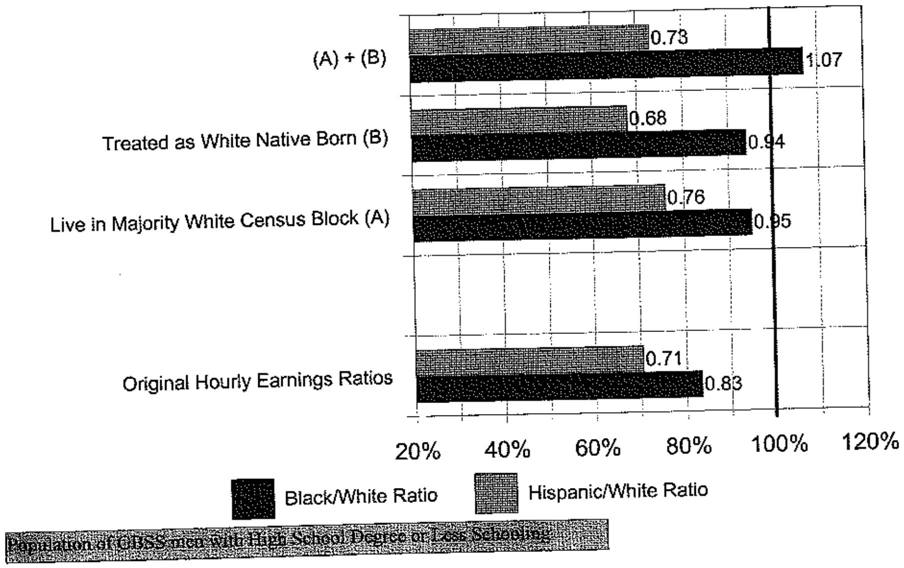
Something well beyond education, years of occupational experience, and improved health is necessary to make further inroads into the large remaining earnings deficit experienced by black men whose formal education ends with a high-school degree. Raising human capital levels to white averages improves black annual earnings by about $2,000. The same exercise raises Hispanic earnings by more than four times that amount — by $8,800. After equalizing human capital, black men still dominate Hispanic men very slightly in hourly wages — but they fall well below them in what really counts — annual earnings.

Race and Residence

If improved human capital does not close the black/white earnings gap very much, what does? Clearly, one possibility might involve ending any racial discrimination and racial segregation. As one test of the efficacy of such a strategy, we simulated the impact of treating both blacks and Hispanics as "native born whites" and "moving" them into white neighborhoods. This is done by setting the variables *BLACK, HISPANIC* and *STRAT* to zero in the logit and regression equations. The results for hourly wages are found in Figure 3.

"Moving" black men into white neighborhoods improves the black/white wage ratio to .95 from its original .83. Similarly, treating blacks as though they were white native born (regardless of where they live) brings the ratio to .94. Doing both simultaneously actually reverses the hourly wage ratio so that blacks dominate whites. This is because in the hourly wage regression, blacks have several factors in their favor including higher unionization rates, a veteran's wage premium advantage, and the advantage of a higher proportion foreign born. Taking these into account, we esti-

FIGURE 3
Men: Hourly Earning Ratios
Impact of Race and Residence



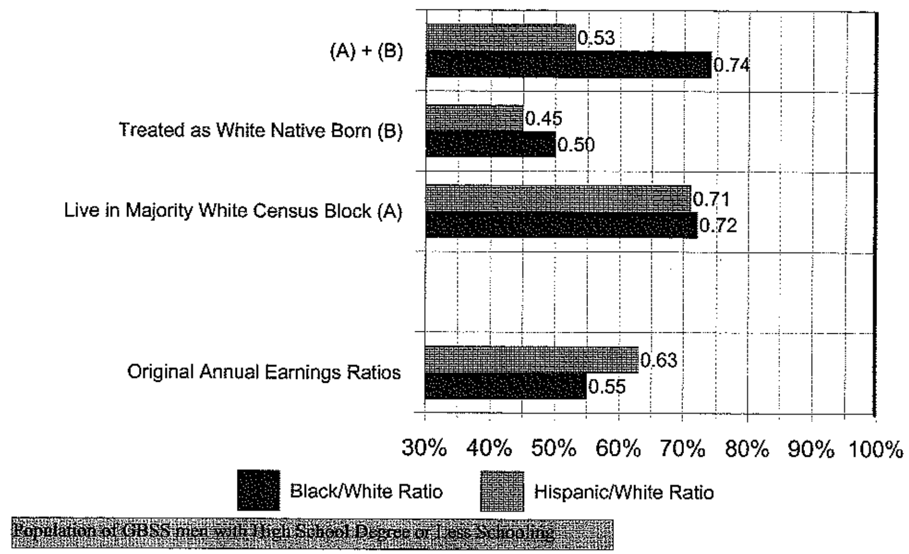
mate that blacks on average would earn 7 percent more than whites — if they were "seen" as white and lived in white neighborhoods. Both race and neighborhood are therefore critical to understanding the existing wage deficit that black men with limited education face.

Hispanic men, on the other hand, improve their position relative to white men by 5 percentage points when they are "moved" into white neighborhoods, but are worse off when treated as white native born. Combining the two effects raises the Hispanic/white hourly wage ratio by only 2 percentage points from .71 to .73. Clearly, for Hispanics, the major problem is a human capital deficit, not discrimination or residency *per se*. In this important aspect, these two minority groups face very different barriers to wage parity with whites.

Turning to expected annual earnings tells a related story — although the outcome for black men is not anywhere near as sanguine as it was in the case of hourly wages. Since black and Hispanic men with limited schooling have measured labor force participation rates that are higher than that of whites, "making" a black or Hispanic into a white actually reduces labor force participation and cuts overall hours. This naturally has a negative impact on expected annual earnings. On the other hand, moving blacks and Hispanics into white neighborhoods improves the earnings ratios considerably. Setting *STRAT* to zero adds 7 percentage points to the black/white ratio and 8 points to the Hispanic/white ratio. (See Figure 4)

Even when we treat blacks as whites and have them live in white neighborhoods, they earn 26 percent less than whites on an annual basis. Ending "discrimination" in this way does more to close the earnings gap than improvements in human capital, but the gap remains very large. For Hispanic men, the annual earnings gap actually

FIGURE 4
Men: Annual Earnings Ratios
Impact of Race and Residence

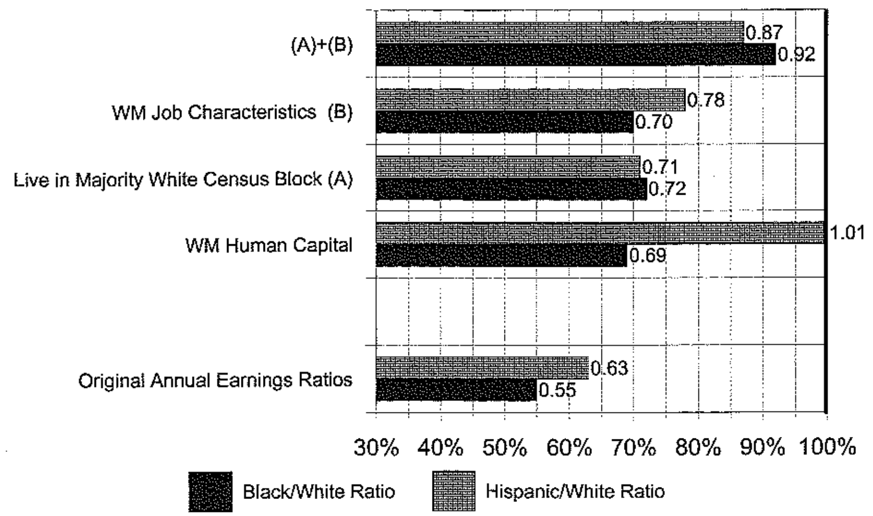


increases when we run this simulation. Treating blacks as whites and placing them in white neighborhoods adds more than $4,000 to their annual earnings. Doing the same for Hispanics drops their annual pay by nearly $2,000. Again, discrimination matters for blacks; human capital deficits matter more for Hispanics.

What *could* close the black/white gap? To help answer this question, we ran an additional simulation in which all we changed were the job characteristics of black and Hispanic men. In this case, we gave the two minority groups the same average number of jobs as their white colleagues (*NUMJOBS*), the same number of years of occupational experience (*TOTJEXPA*), the same proportions in sales and service occupations (*SALES*; *SERVICE*), and the same percentage using computers regularly on the job. All of these variables appear in the weekly hours regression.

The result of this simulation is found in Figure 5. The black/white annual earnings ratio rises to .70 from .55 while the Hispanic/white ratio rises to .78 from .63. In the same chart, we have added the simulation for human capital and for living in a white majority neighborhood. Giving blacks the same job characteristics as whites does more for their annual earnings — by boosting their working time — than does equating human capital. Moreover, combining two simulations, giving blacks both the same job characteristics as whites and moving them into white neighborhoods [(A)+(B)], closes the annual earnings ratio to .92 from the original ratio of .55. This means that four-fifths (80 percent) of the annual earnings gap can be closed in this way (.37/.45), while equalizing human capital closes less than one-third (31 percent) of the original gap. For Hispanics, we close more of the annual earnings gap through human capital than by relocating them to white neighborhoods or by giving them the job characteristics of white workers. By placing blacks in the kinds of neighborhoods

FIGURE 5
Men: Annual Earnings Ratios
Impact of Human Capital, Residence, Job Types



Population of GBSS men with High School Degree or Less Schooling

and jobs that whites occupy, their average annual earnings rise by more than $8,500. Doing the same for Hispanic men raises their yearly pay by more than $5,500.

SUMMARY AND CONCLUSIONS

What we have found in this study is that a tight labor market like that found in the Greater Boston metropolitan area makes it possible for blacks and Hispanics with limited schooling to participate in the labor force at rates equal to that of white men. Moreover, the probability of working some time during the year is equally as high as well. There is, therefore, no "jobless ghetto."

Nonetheless, non-college black men — even after controlling for differences in human capital — earn little more than two-thirds (69 percent) as much as white men. The large gap in expected annual earnings is due to higher unemployment rates and the higher probability of working in part-time jobs with little opportunity for overtime. In contrast, after controlling for human capital, Hispanic men close the annual earnings gap with white men as they have unemployment rates no higher than white men and apparently are not limited in obtaining jobs which permit many hours of work. Although their median hourly wage rate is only 85 percent as high as that of black men, their expected annual earnings are 16 percent higher — based on working 36 percent more hours a year.

We found that annual earnings depend on a whole range of factors — and those important for improving the labor market fortunes of blacks are quite different than those for Hispanics. If we were asked what one set of policies would most improve the

lot of Hispanic men with limited schooling, our answer would be injections of more human capital. If asked the same question regarding non-college black men, we would suggest focusing much more on dealing with what appear to be persistent forms of discrimination in the labor market. The discrepancy in their annual earnings is not so much a function of no experience in the labor market or even lower wages when work is found. It is rather an inability to procure the types of employment that permit black men in this tight labor market to secure steady full-time jobs with the opportunity for overtime and moonlighting if they so choose. Working time has become the key problem for blacks in Boston, not low labor force participation or a total lack of job experience. Greater Boston does not have jobless ghettos. It has a severe mismatch between black workers with work experience needing more work and employers who may not know that there is a ready and willing labor force who could benefit from more job opportunity.

APPENDIX A
PROTOCOL FOR ESTIMATING ANNUAL EARNINGS MODEL

Because of the immense number of variables that theoretically could enter the model — and the wide array of demographic information available in the GBSS — it was necessary to develop a protocol for entering variables into each equation. After much experimentation, the following rules were developed:

(1) All variables included in the model had to represent factors highlighted by economic and sociological theory. These included variables related to human capital, nativity and family status, and job characteristics.
(2) The data set was restricted to men age 21-65 who had no more than 12 years of formal schooling.
(3) The original equations were run on *unweighted* data with four variables which remain in the final equations regardless of their statistical significance. These account for race (*BLACK*), for ethnic group (*HISPANIC*), for neighborhood racial and ethnic composition (*STRAT*), and for high-school completion (*HSDEGREE*). Our focus on race and ethnicity, on neighborhood effects, and on human capital provides the rationale for entering these variables in each equation.
(4) Additional variables were then tested in each model. These include a wide array of human capital and demographic factors as well as variables related to the number and type of jobs held by working individuals. These were retained in the model if they met the 10 percent confidence level. Given the relatively small sample sizes for individual equations, this level of significance was deemed reasonable in order to reduce the possibility of "Type II" errors — rejecting valid hypotheses when they are true [Kennedy, 1994].
(5) The preference for using unweighted data in the estimated equations is based on a careful consideration of the tradeoff between inefficiency due to weighting vs. bias due to not weighting [Graubard and Korn, 1995]. In a sample such as the GBSS where some cases have very large weights relative to others, using weighted data for econometric analysis can introduce enormous heteroskedasticity and therefore extremely inefficient parameters. On the other hand, the bias due to not weighting is a result of the stratification of the sample. By including stratification variables in each equation and attempting to specify each equation as carefully as possible in line with theoretical considerations, the bias due to using unweighted data should be less detrimental than the inefficiency introduced by using weighted data.
[6) Logit equations for labor force participation and unemployment rates were reestimated with weighted data because the unweighted specification produced coefficients that yielded highly unstable simulations.

While the original labor force participation and unemployment rate logit equations normally had the expected signs, counterfactual simulations based on the estimated coefficients yielded participation and unemployment rates which were overly sensitive to small changes in the simulated level of their explanatory variables. This turned out to be due to an interaction amng three factors: the relatively small sample size, the use of a stratified, clustered sampling design for the survey, and the fact that the variables to be explained had mean values close to 1 (participation rate) and close to 0 (unemployment rate). (In related research on women where the participation rates were close to .6, this problem did not appear.) After a series of experiments with different forms of the equation, it was found that a model based on sample weighted data proved substantially more serviceable in the simulation model. In the unemployment equation, we also found that adding the *HSDEGREE* variable to the model generated instability as well. In this one case, the variable was dropped.

NOTES

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1. In our population weighted GBSS sample, over 80 percent of black men with no more than a high-school education live in neighborhoods where a majority of the households are black or Hispanic. Only 34 percent live in neighborhoods where at least 25 percent of the households are officially poor. Only 8 percent are living where the poverty rate equals or exceeds 40 percent.
2. Of the white men in our sample, 19.6 percent report working two or more jobs. However, only 2.7 percent of black men and 3.6 percent of Hispanic men are "moonlighters."
3. In generating expected annual earnings for each demographic group we study, we chose *median* values for hourly wages instead of the more conventional *mean*. This was done in order to minimize the impact of outlying values in the reported GBSS data. Given the relatively small sample sizes for various subgroups and the weighting procedure used in the GBSS stratified, clustered survey frame, it is possible that a few cases with outlying values *and* particularly high weights can have a distorting effect on mean values. By using medians for hourly wages, we reduce this potential distortion.
4. In addition to these variables, many others were created from the GBSS survey and introduced into the equation set. These included such factors as years living in Boston, religion, prison record, firm size, percent minority in establishment, public sector employment, and on-the-job training. While we might have expected these to prove statistically significant based on existing literature and economic theory, they did not. Hence, they too were excluded from the final equations used in the simulation.
5. As noted in Appendix A, we were forced to rely on weighted data to estimate the equations for labor force participation and the unemployment rate. This was necessitated by the poor simulation model behavior of unweighted equations for these two expected annual earnings components. When we tried to use unweighted equation coefficients in the simulation model, we found that the simulation model became severely unstable. For example, turning a black worker into a "white" worker could send the simulated labor force participation rate plummeting toward zero.
6. The simulated mean labor force participation rates for single and married men are as follows:

LFP Rates:	Single	Married
Black men	.80	.90
Hispanic men	.77	.88
White men	.70	.86

7. We expected that education and age would be statistically significant in this equation. Neither was. Further analysis provided some insight as to why this was true. If we restrict the sample to those who are healthy (*HEALTHDY*=0), both *AGE* and *AGESQ* have the expected signs and are statistically significant. Health condition so dominates age, however, that it remains the significant variable in equations in which *HEALTHDY* is not restricted. On the other hand, education was never

statistically significant in any of our runs, even when the sample was not truncated on this variable nor when alternative measures of education were specified. Education is a critical variable in other labor market components, but not labor force participation.

8. Given that the population weighted mean unemployment rate for less-educated black men is more than double the rates for Hispanics and whites, we naturally expected the coefficient on the race variable to be positive, large, and statistically significant. While nominally positive, it and its t-statistic were close to zero. What could have caused this peculiar finding? It turns out that the pattern of coefficients in this equation can be explained by comparing weighted and unweighted data. Recall that the weighted unemployment rates for blacks, Hispanics, and whites are 15.2, 7.1 and 7.1 percent respectively. However, given the GBSS oversample in poorer minority neighborhoods, the sample (unweighted) unemployment rates were quite different. The highest unweighted rate was among Hispanics at 20 percent followed by blacks and whites at 12.7 and 11.6 percent, respectively.

9. For the weekly hours equation, there are valid data on the dependent variable for 245 out of 252 current labor force participants. Of these, 212 were currently employed and 33 currently unemployed. In addition, there are valid data on weekly hours in 34 cases where the respondent worked at some time within the past five years of the GBSS, but was not at work when interviewed. Similarly, for the hourly wage equation, we have valid data on the dependent variable for 228 of the 252 current labor force participants. Of these, 197 were currently employed while 31 were unemployed. In addition, we have valid data on hourly wages in 23 cases where the respondent worked at some time in the previous five years, but was either out of the labor force or unemployed at the time of the survey. Among those currently in the labor force, we have valid weekly hours data for 245, but valid wage data for only 228 because 17 respondents did not provide data from which an hourly wage could be calculated.

10. An exception is found, however, for the largest group of immigrants: foreign born Hispanics. They average about seven hours *less* work per week than native born Hispanics (*FORBORN*: −.746) + *FORBORN*×*HISPANIC*: (−5.92)).

11. The percentage difference in the hourly wage rate is calculated according to the following algorithm: percent difference in hourly wage = $(e^{(b\times 1)} - 1)\times 100$ where b is the regression coefficient.

12. The assessment of English speaking ability did not prove statistically significant in any of the equations. This could be due to the fact that within this sample restricted to those with a high-school degree or less, many may be working in ethnic enclaves or in occupations which do not require great facility with the English language. The overall ability to understand the questionnaire, on the other hand, may reflect a broader indication of competence.

13. Overall, the difference between white and black men on this variable is .89 units. This translates, according to the hourly wage regression, into a wage differential of about .01 log points.

14. In fact, analysis of the interviewer assessment of respondent understanding of the questionnaire (*ASSESSUQ*) indicates that white interviewers of black respondents tended to rate the ability of these respondents somewhat *higher* than the average rating given black respondents by black interviewers. On the *ASSESSUQ* scale of 1= excellent to 5 = poor, the average score for black men with no more than the high-school degree was 2.35 in the cases of white interviewers; 2.65 in the cases of black interviewers. Hence, the point difference is small and there does not appear to be any tendency for white interviewers to downgrade the assessment of blacks. White interviewers tended to score Hispanic respondents a little bit lower in ability than Hispanic interviewers, but again the differences were not large. The average *ASSESSUQ* across all Hispanic male respondents with no more than high school was 2.90. The average score for respondents interviewed by whites was 3.17.

15. We found a substantial difference between whites and the two minority groups when it came to their test scores on the survey's word recognition test. Whites scored an average of 3.84 correct answers out of a possible 7. Blacks and Hispanics scored 2.28 and 2.30 respectively. Difficulty with English was not a factor in this test since the test was given with a set of Spanish words for those Latinos who wished to conduct the entire interview in that language. Unlike the interviewer assessments, the word test is not subject to interviewer bias. That the results of the word test are correlated with the interviewer assessment rankings, we find additional reason for having confidence in the survey's assessment questions.

16. Moreover, to test for the possibility of interaction effects masking the impact of high-school completion, we generated interaction terms for *HSDEGREE* and *AGE*, and *HSDEGREE* and a dummy variable for employed in manufacturing. However, neither of these interaction terms proved statistically significant, nor did they materially raise the t-statistic on the education variable. We are forced to conclude that in this truncated sample, the high-school degree *per se* is not a key variable in the determination of labor market success in this particular labor market.

17. According to an analysis of pay data, Lawrence Mishel, Jared Bernstein and John Schmitt [1999] found that the union hourly pay premium in 1997 was 23.2 percent for all workers. It was especially large among blue-collar workers (50 percent).

18. This positive relationship is consistent with national studies. Using 1989 national data and a wage equation not unlike ours, Princeton economist Alan Krueger [1993] finds that the use of a computer at work is associated with an 18.8 percent wage differential. In a more recent paper, John E. DiNardo and Jörn Steffen-Pischke [1996] have challenged the conventional wisdom regarding the importance of computers in terms of wage and productivity premia. Using German data they have shown that computer use is indeed associated with higher wages. But so is the use of pencils, calculators, and telephones! What they argue is that it is not the computer use *per se* that contributes to higher productivity and pay, but the jobs that require any kind of higher-level technical and social skills.

19. Several adjustments were needed in the model in order to align our estimates with the population weighted means for each labor market outcome component. After much experimentation, we chose the following adjustments as most appropriate.
    (1) After entering all logit or regression coefficients for each component in the baseline spreadsheet and calculating mean outcomes for black, Hispanic, white, and all men, the constant in the fitted equation for each component was adjusted to bring the fitted mean value for *all men* into alignment with the actual mean value.
    (2) After this was completed, any difference between the fitted mean value and the actual mean value for *each race/ethnic group* in the baseline spreadsheet was adjusted to zero by adding or subtracting to the fitted value until it was aligned with the mean value.

This process permits us to begin the simulation with race/ethnic group "fitted" means equal to the actual sample weighted means for each race/ethnic group. Moreover, this adjustment process permits the values estimated in the various simulations to reflect differences from the actual weighted baseline means rather than differences from the means estimated on unweighted data.

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