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GENERAL & APPLIED ECONOMICS | RESEARCH ARTICLE

Latent earning capacity and the race marriage gap

Robert Nakosteen^{1*} and Michael Zimmer²

Abstract: One of the most persistent socioeconomic phenomena in the process of family formation is the relatively low rate of marriage by black men and women. The enduring conventional wisdom has been that low black marriage rates reflect a relative shortage of marriageable black men. Yet numerous studies that have attempted to account for the shortage have reported that the race marriage gap remains, albeit sometimes in reduced magnitude, even after controlling for economic attributes of potential spouses and potential supplies of spouses in regional marriage markets. This paper examines the possibility that the race gap is explained in part by disparities in unobserved earning capacities between black and white men. In doing so, this study redefines the marriage market for each man in terms of his position in the earnings distribution rather than by geographic region. Our results indicate that when young black men are placed in competitive positions in the distribution of white residual earnings, the race gap disappears and even shows strong signs of greater marriage propensities in the black population.

Subjects: Microeconomics; Labour Economics; Econometrics;

Keywords: race; marriage; earnings distribution; latent earnings

ABOUT THE AUTHOR

Professors Robert Nakosteen and Michael Zimmer have been studying labor market outcomes, and their relationship to individual decision-making, for nearly 40 years. Their focus has been on the influence of self-selection and latent or unobservable variables. Decisions such as whether to move, to marry, to divorce, or to participate in the labor market have been researched. These decisions are modeled in the context of their effect on labor market earnings, or the effect of labor market earnings on the decision. Their research has been published in a variety of academic journals, including the *Journal of Regional Science*, the *Journal of Population Economics*, *Papers in Regional Science*, *Economic Inquiry*, and the *Journal of Human Resources*, among others. The interdependent relationship among race, earnings, and marriage rates is an extension of this research program.

PUBLIC INTEREST STATEMENT

Social scientists recognize that the incidence of formal marriage in recent decades is lower among blacks than among other race and ethnic groups. Research has tended to attribute the disparity to relative scarcities of marriageable black men in “marriage markets” for unmarried black women. Others have cited a general propensity to refrain from marriage in the black population. This paper adopts a different approach, addressing the disparity as a deficiency in earnings among black men relative to other race groups. Using U.S. population data, we employ a statistical approach that synthetically imputes to black men latent capacities for earnings that are comparable to white men. When endowed with those hypothetical earning capacities, black males display propensities for marriage that are similar to (and in some cases exceed) those of white men. The implication is that the deficiency of marriage in the black population is more a manifestation of comparatively low earnings among black men than it is a reflection of culture or race *per se*.

JEL classification: J12; J15

1. Introduction

The United States has experienced a persistently low rate of marriage in the black population relative to the white population, and the marriage race gap appears to have increased over time. For 2015, data from the Current Population Survey indicate that, among white men, the proportion married was 0.56. The comparable proportion for 1970 was 0.68. Among black men, the proportion married in 2015 was 0.38, reflecting a decline from 0.57 in 1970.¹

An important social challenge attendant to the marriage race gap is the incidence of child-bearing out of wedlock among black females (see, for example, Olsen & Farkas, 1990, Figure 1; Keane & Wolpin, 2010, Table 1). The resulting prevalence of households in the black population headed by never-married mothers is well known to economists and policymakers.

Numerous contributions to the literature have attempted to account for the marriage race gap. A widely discussed hypothesis is Wilson (1987), who attributes the gap to inadequate supply of marriageable males in the black population. An elaboration of that argument advanced by Wilson and Neckerman (1986, p. 253) is that, in particular, “... the increasing rate of joblessness among black men merits serious consideration as a major underlying factor in the rise of single black mothers and female headed households.” Banks and Gatlin (2005, p. 126), however, argue that the retreat from marriage among black men extends to those who attain higher levels of education, and presumably have greater prospects for remunerative employment.

There is a puzzling feature, however, about the rather sizable literature that has arisen subsequent to the Wilson and Neckerman (1986) hypothesis. The tendency among those studies, to be briefly surveyed in the following section, is to conclude that the shortage of marriageable black males explains at best only a portion of the marriage race gap. In particular, a common finding is that a significant race gap persists even in models of marital formation that control for the available stock of potential husbands.

This paper extends the literature in two respects. First, in changing emphasis from the supply of marriageable men to their economic suitability as potential husbands, it focuses on latent capacities for earnings and how those capacities explain transitions by single men into marriage. Second, we depart from the convention of defining a man’s marriage market in terms of geographic region. Instead, we identify each man’s “location” in terms of his position in the distribution of male earnings. The implied premise is that single men who are favorably located in the distribution of earnings are relatively attractive as marriage partners.

Analysis in this study is based on a pooled series of two-period quasi panels constructed from successive years of the March Current Population Survey (CPS). By exploiting the outgoing rotation groups in the CPS, we are able to collect samples of young white and black men who are never married in the first year of each two-year sequence. We then use the second year outcome of marital status to observe whether each man marries or remains single. Our approach entails the delineation of earnings-based partitions within the sample. Treating those groups as distinct marriage markets, we estimate marriage race gaps within each market.

Our results are noteworthy. The race gap declines substantially when we quantify the individuals’ latent earning capacities and account for their locations in the earnings distribution. Indeed, the estimates indicate that black males who are well positioned, in the statistical sense, are more likely to marry than otherwise comparable whites. We interpret that finding as complementary to the Wilson and Neckerman (1986) hypothesis: because of the shortage of marriageable black men,

Table 1. Marital status by race: proportion married year 1 and year 2

Age range	White		Black	
	Year 1	Year 2	Year 1	Year 2
18–29	0.33	0.35	0.13	0.17
30–40	0.74	0.71	0.45	0.45
N	28,412		3,698	

Year 1 refers to the first year of the two-year sequence for each man: 2002, 2003, 2004, 2005, 2006, 2007, 2008, 2009. Year 2 refers to the follow-up year for each initial year: 2003, 2004, 2005, 2006, 2007, 2008, 2009, 2010.

the marriage market places a race premium on the highest potential earners in the black population.

2. Literature review: the race marriage gap

Early work by Wilson and Neckerman (1986) and Wilson (1987) suggested that the race marriage gap is the result of deficiency in the supply of marriageable males within the black population, which is compounded by low levels of employment among black males of marriage age. Subsequent researchers examined that thesis, using a variety of data sets and empirical methodologies.²

Schoen and Kluegel (1998) find only weak support for the Wilson and Neckerman (1986) hypothesis. Using regionally aggregated data, they instead attribute the race gap to a general decline in propensities of black females toward marriage. Wood (1995), using data aggregated at the level of Standard Metropolitan Statistical Areas (SMSA), likewise reports that the Wilson and Neckerman (1986) hypothesis offers only marginal explanation for the proportion of ever-married populations across SMSAs.

Olsen and Farkas (1990) use experimental data to model transitions to marriage among black females. They find relatively weak support for the Wilson and Neckerman (1986) hypothesis. Instead, their evidence points to a strong role for economic opportunities in dissuading females from bearing children out of wedlock and inducing them to enter cohabiting or married unions.

Other research exploits panel data to examine transitions to marriage. Lichter, McLaughlin, Kephart, and Landry (1992) analyze marital transitions for young women. Their results indicate that a significant race gap among women remains even after controlling for the supply of marriageable men in local marriage markets. The implication is that the Wilson and Neckerman (1986) hypothesis is not sufficient to explain the race gap. It is noteworthy for the purpose of this study, however, that Lichter et al. (1992) report a positive propensity to marry among higher educated and employed women and those with higher earnings.

Brien (1997) applies a discrete time logit model of marriage transition for females, while Raley (1996) uses a proportional hazard model for time to the formation of cohabitation unions or marriages by females. Both studies conclude that the marriage market hypothesis of Wilson and Neckerman (1986) offers only a partial explanation of the marriage race gap. McLaughlin and Lichter (1997) likewise model transitions to marriage for women. Similar to the studies cited above, they conclude that marriage market factors, in particular, the supply of marriageable men, do not offer a complete explanation of the race gap. Focusing on the poverty status of women, the authors report a result that is indirectly relevant to the present study, namely that black women in poverty exhibit marriage propensities similar to white women in poverty.

Bennett, Bloom and Craig (1989) attribute the marriage gap to declining trends in relative earnings and employment among black men. In a point related to the present study, they note a relatively strong propensity to marry among higher educated black women.

Focusing on transitions to marriage among young single men, Lloyd and South (1996) use individual panel data with region-based variables measuring the ratio of potential female spouses to young males. Treating individual income as an exogenous variable, they infer that high-earning males of both races are more prone to marry. Estimates from a pooled sample of blacks and whites indicate that the race gap persists even after controlling for the local ratio of marriageable women to men.

The first marriage gap study of which we are aware that explicitly accounts for male earnings as an endogenous variable in a structural model of marital status is Nakosteen and Zimmer (1991). They argue that the race gap is illusory; after correcting for earnings, black males appear to marry at a rate that does not differ significantly from whites. However, their model is a static representation of marital status at a point in time and consequently does not address transitions into marriage.

Recent research has attempted a more detailed accounting of the race gap. Using U.S. Census data, Charles and Luoh (2010) identify male incarceration rates as significant in determining marriage rates of young women. Seitz (2009) and Keane and Wolpin (2010) examine the role of individual earnings in explaining the marriage gap. Keane and Wolpin (2010) develop a structural model of life cycle decisions for young women, including marriage. In their model, an experiment that synthetically equalizes wages of young black and white men has the effect of increasing the black marriage rate nearly to parity with whites. Seitz (2009) likewise conducts a counterfactual experiment involving wages in the context of a structural model of employment and marriage. Her outcome contrasts with that of Keane and Wolpin (2010). Removing the gender wage gap has the consequence of *increasing* the marriage gap between blacks and whites.

The studies surveyed here, using a variety of data sources and statistical methodologies, seem to produce two broad conclusions. First, evidence in support of the Wilson and Neckerman (1986) hypothesis is mixed, and in general somewhat weak. Specifically, controlling for marriage markets in a geographic context and using various constructs to measure demographic ratios of available spouses, the white-black marriage gap persists in the majority of studies. Second, it appears that earnings of single men and women play significant roles in transitions to marriage. Using measured earnings as the variable of choice, most studies report propensities to marry for high earners of both genders.

This paper extends the literature in both respects. In particular, we attempt to unify the marriage market and earnings phenomena. As introduced in Section 1, and focusing on young single men, we define each individual's marriage market in terms of his location in the earnings distribution rather than by geographic location. In addition, we identify the earnings location on the basis of latent earning capacity, i.e., unmeasured skill and earning attributes, rather than by measured earnings. The following section formulates a statistical framework for this approach.

3. Statistical framework

3.1. A model for latent earnings and the transition to marriage

Our empirical approach focuses first on earnings of young never married black and white men. The nature of the data permits us to observe each man for two consecutive years.³ In year one, we measure the logarithm of his wage earnings from labor, denoted Y_{i1} , where i indexes the individual, along with a set of individual characteristics and background variables that potentially determine earnings. Adapting the framework of Juhn, Murphy, and Pierce (1991), first period earnings are expressed as

$$Y_{i1} = \beta' x_{i1} + \sigma \theta_{i1} \quad (1)$$

where x_i denotes a vector of observed explanatory variables and β is a conformable vector of unknown coefficients. The unmeasured or latent component of earnings is represented by the second term, where θ_{i1} is a random error term distributed as independent normal with zero

mean and unit variance. In the terminology of Juhn et al. (1991, 1993), β and σ represent skill prices for measured and unmeasured characteristics, respectively.

Previous researchers have relied on *measured* earnings or income as explanatory factors in determining transitions of single men into marriage (see, for example, Lloyd & South, 1996). In this study, we focus on unmeasured (to the econometrician) characteristics that affect observed earnings and, in turn, marriage transitions. These characteristics, while not captured in conventional data sets, are likely assessed by potential spouses of young men. They include personality traits such as maturity and stability, ambitions and aspirations, plans for future human capital investment, propensities for risk-taking, and perhaps even family connections that portend future potential for earnings or wealth.

Young men who are endowed with favorable latent earning capacity exhibit a positive disturbance in the θ_{11} term of the earnings equation. Accordingly, their earnings exceed the regression-adjusted population mean in a manner that appears (to the econometrician, owing to limitations in most data sets) to be random and unexplained by the model. It is likely, however, that those disparities are not completely unexplained to the potential spouses who make their acquaintances. Indeed, the likely scenario is that potential spouses actively search for young men who populate the most favorable percentiles of the earnings distribution. This is the basis for the conception of marriage markets in the present paper. Our approach is to partition marriage markets based on the distribution of earnings for unmarried men. The premise is that men who populate the high end of the distribution face a market of potential spouses that is distinct from those at the low end. The key point to emphasize is the latent nature of this phenomenon. It most likely cannot be captured as a directly measured characteristic, and yet it is certainly known to the men themselves and to the pool of potential wives.

As noted above, we have a second year of data for each man, in which we observe whether or not he transitions to marriage. Denote M_{i2} as the individual's propensity to marry in period two. He chooses to marry if $M_{i2}^* > 0$; otherwise he remains single during the second period. The marriage propensity is a linear combination of explanatory variables and a random error term:

$$M_{i2}^* = \delta' z_{i1} - \omega_i \quad (2)$$

where z_{i1} represents a set of characteristics observed in the first period and δ is a conformable vector of unknown coefficients.

The marriage propensity is not observed. Instead, we observe a dichotomous variable

$$M_{i2} = 1 \text{ if } M_{i2}^* > 0, \text{ and} \quad (3)$$

$$M_{i2} = 0 \text{ if } M_{i2}^* \leq 0.$$

Subject to the assumption that the random error term in (2) is normally distributed with zero mean and unit variance, the probability of married status in period two is given by

$$P(M_{i2} = 1) = \Phi(\delta' z_{i1}), \quad (4)$$

where Φ denotes the cumulative distribution function associated with the standard normal distribution. Accordingly, the parameters in (4) can be estimated by means of maximum likelihood for a probit model.

Included among the explanatory variables in (4) is a dummy variable identifying black men, with white as the reference category. It serves as a measure of the white-black difference in males' transitions to marriage.

3.2. Empirical approach

Our empirical approach begins with the initial-period earnings Equation (1), with particular focus on unmeasured skills as represented in the second term. In order to normalize the distribution of latent skills between races, the first step is to estimate the earnings equation for white men in period one. The resulting estimated parameters are $\hat{\beta}_W$ and $\hat{\sigma}_W$, where the subscripts denote that the estimates are based on the white sample.

In step two, we obtain an estimate of latent earning capacity for each black man in period one, using the estimated skill prices for white men from the first-step regression:

$$\hat{\eta}_B = \hat{\sigma}_W \hat{\theta}_B = Y_{i1,B} - \hat{\beta}_W x_{i1,B} \quad (5)$$

These estimates are counterfactual in the sense that they attempt to capture each black man's unmeasured earning capacity under the restriction that his skill prices are determined in accordance with the white distribution.⁴

We next estimate earning capacity for each white man:

$$\hat{\eta}_W = \hat{\sigma}_W \hat{\theta}_W = Y_{i1,W} - \hat{\beta}_W x_{i1,W} \quad (6)$$

In the empirical analysis, we use this approach to construct a statistical marriage market for each black man and pool it with a subset of white men who serve as a control group. To illustrate this idea, Figure 1 imagines a marriage market comprising high skilled single men of both races. The first graph represents the distribution of standard earning residuals for whites. The shaded area identifies white men whose values of latent skills place them above the white population mean. The bottom graph is drawn, for purposes of illustration, to show a hypothetical scenario in which the population of black latent skills possesses a counterfactual mean, controlling for white skill prices in both measured and latent characteristics, that is positioned to the left of the white mean.⁵ The shaded area in the second graph identifies a subset of black men whose latent skills are sufficiently high valued to place them at or above the median of the white population.

This describes one of several scenarios that we construct in Section 5. In each case, we combine the two subsets of interest into a single sample for purposes of estimating the marriage transition Equation (4). The objective is to estimate the race gap, if any, that exists in each scenario. The motivating idea of this paper is that the race gap, which has been widely and persistently reported in the literature (see Section 2), might be in part an artifact of unmeasured earnings differences rather than a result of race *per se*.⁶ Taking for example the scenario illustrated in Figure 1, if the phenomenon is in part due to a latent earnings gap, then we should not see a black versus white deficiency in transition to marriage in a synthetic marriage market characterized by and restricted to the shaded areas, an empirical approach that holds constant the distribution of latent skills.

4. Data and model specification

4.1. The current population survey and outgoing rotation groups

Estimation of the model is based on data from the United States Current Population Survey (CPS) for 2002 through 2010. The annual March CPS contains data from approximately 50,000 U.S. households. A selected sample of households is interviewed for four consecutive months, before leaving the sample. These outgoing rotation groups are then omitted from the survey for eight months, after which they return for another four-month period.

Since CPS samples are based on addresses rather than individuals or households, we took steps to ensure to the greatest extent possible that the resulting data constitute legitimate quasi-panels for each man. Across the two years, we matched gender, race, age and education. We deleted observations that did not show consistency between the incoming and outgoing rotations. Thus, although the core content of the CPS is essentially cross-sectional, the rotation groups allow us to

access two adjacent years of data for each individual. This is necessary for estimation of the model represented in Section 3, since it purports to explain the individual's transition to marriage, if any, in year t based on his exogenous controls in year $t-1$. For the first outgoing rotation group, the base year is 2002 and the year for potential marriage is 2003. For this group, the only variable extracted from the 2003 CPS is M_{iz} in Equation (4), the dichotomous indicator denoting marriage. All other variables in the model are taken from the base year. Similarly, for the second outgoing rotation group, the base year is 2003 and the outcome year is 2004, and so on, up to the final group, with base year 2009 and outcome year 2010. The sample consists of male salary and wage earners, aged 18 to 40, all of whom are never married in their respective base years, are not self-employed and not in the armed forces. The sample used for estimation consists of 7,849 men, of whom 741 married between their base and outcome years. In order to assemble adequate samples for the scenarios described in Section 3, we pooled the eight outgoing rotation groups. Accordingly, as described below, the model deflates the earnings over time to reflect constant (Year 2002) dollars.

In addition to individual earnings and marital status outcomes, the data file contains background information for each individual, including age, education, race, disability status, union membership, and occupation, all of which are included in the earnings equation. Included exclusively in the marriage outcome equation is a dummy variable indicating home ownership as an indicator of wealth. In addition, since the sample pools outgoing rotation groups from 2002–2003 through 2009–2010, the model includes dummy variables in the marriage outcome equation as controls for the individual cohorts. The cohort effects attempt to capture pervasive macroeconomic or other cohort-wide changes that potentially affected marriage formation throughout the sample period.

4.2. Attrition in the current population survey

As noted above, analysis in this study requires strategically chosen subsets from within the population of single men. This sampling scheme requires a large parent sample from which the subsets can be extracted, and the subsets must be quasi-longitudinal in the sense that they include a second period of observation for each individual that reveals his marriage outcome. Data resources that are suitable for this task appear to be quite limited, particularly in the context of the United States. Our choice of the CPS is motivated by its two principal advantages. The first is clearly its large size, with monthly coverage of approximately 50,000 households from all ethnic groups and every geographic region. Second, availability of the outgoing rotation groups makes it possible to construct the necessary short panels for approximately one-eighth of the CPS sample in each year. As Neumark and Kawaguchi (2004) point out, these advantages have been exploited by researchers to study a wide range of phenomena in labor economics.

One challenge in using the CPS is that it does not follow individuals who change their places of residence. Consequently, short panels obtained from the outgoing rotation groups are characterized by year-to-year attrition, estimated by Neumark and Kawaguchi (2004) to be as high as 30%. The extent to which attrition is problematic depends of course on the application for which the data are used and whether those who leave the sample represent a nonrandom subset of the population. In the context of this study, recalling that we use the second period to observe (only) marriage outcomes for single men, the challenge of nonrandom attrition translates to the question of whether men who marry also become residential movers to an extent that imparts bias in our parameter estimates.

This issue is examined in an innovative study by Neumark and Kawaguchi (2004), who examine CPS attrition in the context of the marriage wage premium for men who transition from single to married status. Their approach is to mimic the CPS by constructing a short panel from the Survey of Income and Program Participation (SIPP). The SIPP is well suited for this purpose, first because it uses a sampling frame and design similar to the CPS and, second, because the SIPP undertakes substantial effort to track residential movers. Neumark and Kawaguchi (2004) proceed in two

steps. First, they estimate the marriage premium based on the full SIPP sample. Second, they remove residential movers from the sample, thus simulating CPS attrition, and repeat the estimation. Their descriptive statistics reveal that the incidence of residential migration in conjunction with changes in marital status (married to not married and the reverse) is approximately 13%. Their estimates of the marriage wage premium suggest that using the matched CPS files imparts attrition bias in the direction of underestimating the premium. However, they find the magnitude of bias to be relatively small and conclude (Neumark & Kawaguchi, 2004, p. 470) that for many applications the advantages of using the CPS short panels outweigh the disadvantages from attrition bias. Based on their evidence, we proceed to exploit those advantages with the caveat that the estimates are not likely to be completely free of bias, particularly if there are unmeasured differences by race in the transition to marriage that simultaneously and strongly determine the propensity for residential migration.

4.3. Model specification and descriptive statistics

The model is completed by specifying exogenous variables for inclusion in Equations (1) and (4). In Section 5 we report estimates for the following specifications:

Tables 1–3 provide a descriptive overview of the sample. Table 1 is based on the larger CPS universe of men aged 18 to 40 from which we extracted the working samples for empirical analysis. As described in the previous section, the working sample is restricted to men who are not married in year one of the two-year sequence. For illustrative purposes in Table 1, however, we take *all* men in age groups 18 to 29 and 30 to 40 and estimate the proportion married in each year. In both age groups, the table reveals the lower incidence of marriage among black men that has been commonly reported in the literature. The race gap does not appear to diminish across age groups. For example, the estimated gaps in year 2 are 18 percentage points and 26 percentage points for the 18–29 and 30–40 age groups, respectively.

Table 2 describes the nature of sample selection by race and age group. The first two rows reveal, consistent with Table 1, the greater proportion of unmarried black men relative to white men. Row three calculates the simple marriage gap in Year 1 by race and age group, and

Single status declines by 41 percentage points for white men and 32 points for black men.

Row 4 shows the difference-in-differences. The white-black marriage gap is 39 percentage points among men aged 18 to 29, and 57 percentage points among men aged 30 to 40.

The sample selection process described above is necessitated by basic objective of this study, and it is generally characteristic of other research in transitions to marriage. Examples include

Equation (1)	Equation (4)
Year 1 Earnings (2002 \$)	Year 2 Marriage Outcome
Constant	Constant
Age	Race: Black
Age squared/100	Age
Education	Age squared/100
Race: Black	Own home
Union status	Education
Disability status	Disability status
Professional occupation	Cohort dummy variables: 2003–2009

Bennet et al. (1989), Olsen and Farkas (1990), Lichter et al. (1992), McLaughlin and Lichter (1997), Lloyd and South (1996), and Nakosteen and Zimmer (1997).

Table 3 presents the sample means for the remaining variables in the model. It is based on the restricted sample of men with positive earnings and not married in year one of the two-year sequence. The sample means appear to reflect the select nature of the sample. They indicate that earnings and years of completed education are nearly equal across race groups. Black men are older by approximately two years on average, and they show higher rates of union membership and professional occupation status. On the other hand, the rate of home ownership is lower among black men relative to white men.

5. Results of estimation

5.1. Preliminary estimates

Our empirical approach, described in Section 3, entails a series of scenario-based samples of black and white men that arise from placement of selected black men into the residual earnings distribution of white men. Criteria determining the black men to be selected in each case depend on the particular comparisons of interest in the white distribution. For the example shown in Figure 1, the comparison entails above-median white men with a subset of black men whose (counterfactual) residual earnings place them likewise in the upper half of white residual earnings.

To set the stage for this and other useful comparisons, in Table 4 we present estimates of the earnings Equation (1) and a descriptive probit equation similar to (4). Columns one and two present estimates of the earnings equation in 2002 for never married blacks ($N = 306$) and whites ($N = 1,096$) respectively. The dependent variable is the logarithm of annual earnings from wages in 2002. Column three presents the probit estimates of marital status, estimated for all first-year observations in the sample ($N = 22,495$). The probit equation is presented only for descriptive purposes, in order to illustrate the race gap in marital status. Consequently, unlike the refined sample used in estimation, these estimates are based on all men in year one, both married and single, including those not employed and those who are employed but report zero earnings for the survey period.

The earnings estimates indicate the familiar concave profile with respect to age, a proxy for work experience. The education coefficient indicates an annual return to schooling of approximately 13% for black men and 5% for white men. There is, as expected, a substantial premium to employment in a professional or technical occupation. While the union and disability coefficients are large in magnitude, it is worth recalling the young age of the sample. With respect to the union coefficient, at young ages, the union advantage in early-career earnings is substantial relative to, say, recent college graduates of the same age. As the careers of both groups mature, the union

Table 2. Sample selection and year 1 marital status

Age	White		Black	
	18–29	30–40	18–29	30–40
Unmarried	0.672	0.265	0.867	0.552
Married	0.328	0.735	0.133	0.448
[Unmarried–Married]	0.344	–0.470	0.734	0.104
[Unmarried–Married]—White [Unmarried–Married] Black	–0.390	–0.574	–0.390	–0.574

Year 1 refers to the first year of the two-year sequence for each man: 2002, 2003, 2004, 2005, 2006, 2007, 2008, 2009. Sample selection is based on unmarried men in Year 1.

Table 3. Variable definitions and sample means

Variable	Definition	White	Black
Y ₁ :Equation (1)	Logarithm of individual's annual labor earnings in year one (constant Year 2002 dollars)	9.615	9.661
AGE	Age in Year 1 of the two year sequence	26.3	28.4
EDUC	Years of education	13.4	13.2
UNION	=1 if individual was member of labor union =0 otherwise	0.083	0.108
PROF_OCC	=1 if individual's occupation is professional or managerial =0 otherwise	0.109	0.137
DISAB	=1 if disability limits or prevents work =0 otherwise	0.018	0.014
BLACK	=1 black =0 if white		0.160
OWN_HOME	=1 if individual owns home outright or through mortgage =0 otherwise	0.624	0.464
YEAR1 = 2002	=1 if year one is 2002 =0 otherwise	0.166	0.244
YEAR1 = 2003	=1 if year one is 2003 =0 otherwise	0.128	0.116
YEAR1 = 2004	=1 if year one is 2004 =0 otherwise	0.109	0.110
YEAR1 = 2005	=1 if year one is 2005 =0 otherwise	0.107	0.092
YEAR1 = 2006	=1 if year one is 2006 =0 otherwise	0.117	0.106
YEAR1 = 2007	=1 if year one is 2007 =0 otherwise	0.128	0.112
YEAR1 = 2008	=1 if year one is 2008 =0 otherwise	0.120	0.114
YEAR 1 = 2009	=1 if year one is 2009 =0 otherwise	0.123	0.104

premium at later ages is likely to moderate. Regarding the disability coefficient, it is likely that, at young ages, workers with impairments have yet to migrate to specialties and occupations that efficiently accommodate their most advantageous divisions of labor. As a result, the negative impact of disability is likely to be relatively large at young ages.

Column three presents probit estimates of married status in year one, using the broadly inclusive sample as noted above. As noted above, by design and for illustrative purposes, these estimates are cross-sectional and preliminary in nature, and they do not control for the role of latent earnings. This model does control for age, education, disability status, wealth in the form of home ownership, and cohorts by year. These controls are discussed briefly in the results to follow, but the essential point is that, absent the earnings control, the race gap indicates a strong aversion among black men for marriage. This is consistent with what has been commonly reported in the literature.

Table 4. Preliminary estimates

Variable	Log Annual 2002 Earnings: Black	Log Annual 2002 Earnings: White	Married in Year 1: All Years
CONSTANT	4.534 (4.49)	2.176 (4.41)	−10.278 (39.40)
BLACK			−0.609 (20.24)
AGE	0.204 (2.82)	0.432 (11.77)	0.606 (35.10)
AGESQ/100	−0.003 (2.22)	−0.006 (3.14)	−0.009 (30.54)
DISAB	−1.394 (4.15)	−0.612 (3.14)	−0.471 (5.98)
EDUC	0.133 (6.49)	0.046 (4.87)	−0.008 (2.39)
UNION	0.189 (1.01)	0.167 (1.79)	
PROF_OCC	0.121 (1.22)	0.237 (4.41)	
OWN HOME			0.597 (28.61)
YEAR1 = 2003			−0.071 (2.17)
YEAR1 = 2004			−0.161 (4.61)
YEAR1 = 2005			−0.047 (1.36)
YEAR1 = 2006			−0.111 (3.24)
YEAR1 = 2007			−0.117 (3.43)
YEAR1 = 2008			−0.095 (2.76)
YEAR1 = 2009			−0.146 (4.20)
Adj R square	0.317	0.365	
N	306	1,096	22,495

Figures in parentheses are absolute *t* statistics. Earnings regressions are for positive earners only.

5.2. The marriage race gap: Tables 5–9

As noted previously, Figure 1 illustrates the first of several scenarios that form the bases for estimates of the marriage race gap. The estimates, based on Equation (4), are presented in Tables 5–9. Each table and scenario corresponds to a distinct synthetic marriage market, based on the distribution of latent earning capacity. In the CPS data, some men report themselves as usually employed but with zero earnings for the survey period. Accordingly, in Tables 5–9 we report estimates for the sample of self-reported workers (column one in each table) and a second set of estimates after excluding zero earners (column two).

5.3. Above median whites and right tail blacks

Estimates in Table 5 correspond to the depiction in Figure 1. Accordingly, the sample consists of whites whose latent earnings are above the white median and blacks whose latent capacities would position them above the median if they were placed in the white distribution. This results in a sample of 4,108 men when all usual workers are included and 3,836 when zero earners are excluded. In the latent earnings sense, this subset of unmarried black men is competitive in the marriage market for white men. It is worth repeating that the latent, or unmeasured, nature of their relative earnings strengths arises after controlling for their own measured characteristics and after imputing the prices of measured and unmeasured skills associated with the white distribution.

Estimated coefficients of the control variables suggest a concave age profile with respect to transition to marriage within the range of age (18–40 years) in the sample. Disabled men are significantly less likely to marry, and the cohort dummy coefficients corroborate findings from other studies that suggest general pronounced retreat from marriage during the past decade. (See, for example, Amato (2010) and references therein.) The coefficients for years of education and home ownership are estimated to be negative, but neither of them differs significantly from zero at levels of significance below 10%.

Table 5. Probit estimates: right tail blacks with above median whites

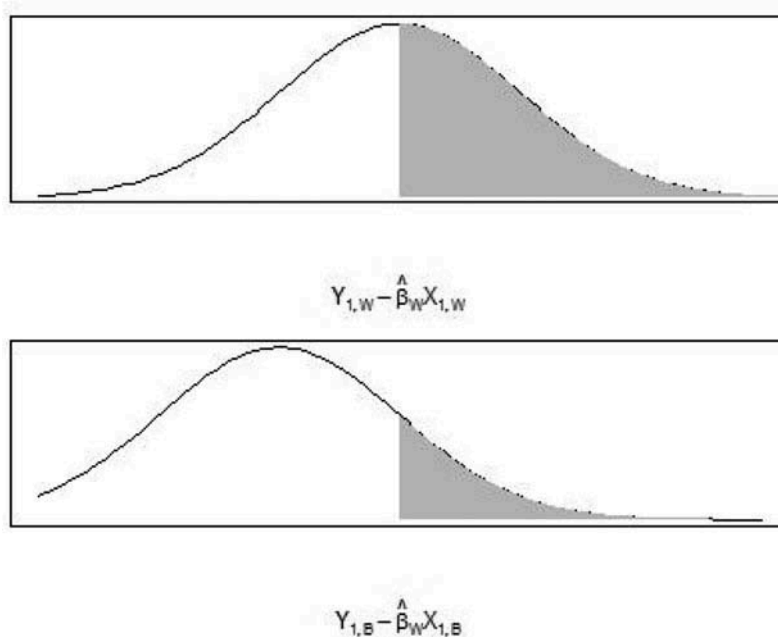
Variable	All self-reported workers	Excluding zero earners
BLACK	0.130 (1.76)	0.093 (1.22)
OWN HOME	−0.025 (0.45)	−0.019 (0.34)
AGE	0.124 (3.00)	0.126 (2.88)
AGESQ/100	− 0.002 (2.65)	−0.002 (2.57)
EDUCATION	−0.014 (1.56)	−0.015 (1.62)
DISABLED	−0.746 (2.40)	−0.734 (2.34)
YEAR1 = 2003	−0.470 (4.75)	−0.427 (4.27)
YEAR1 = 2004	−0.391 (3.90)	−0.404 (3.90)
YEAR1 = 2005	−0.491 (4.50)	−0.423 (3.84)
YEAR1 = 2006	−0.344 (3.49)	−0.345 (3.39)
YEAR1 = 2007	−0.349 (3.66)	−0.334 (3.43)
YEAR1 = 2008	−0.429 (4.26)	−0.390 (3.80)
YEAR1 = 2009	−0.552 (4.75)	−0.517 (4.85)
CONSTANT	−2.650 (4.62)	−2.637 (4.39)
N	4,108	3,836

Figures in parentheses are absolute *t* statistics. Dependent variable = 1 if individual transitioned to marriage, =0 otherwise.

The most striking feature of Table 5 is the reduction in the race gap. Indeed, the estimated coefficient of black status is marginally *positive*, although not strongly significant. In column one, the estimate is 0.13 (*t* = 1.76), which is significant at the 0.08 level. Exclusion of zero earners (column two) reduces the estimate to 0.09 (*t* = 1.22), significant at the 0.22 level.

The implication is that the ostensible race marriage gap is at least in part an earnings phenomenon. When we extract relatively strong latent earners from the black population and place them in a statistical sense with above-median whites, they are not less likely to marry. On the contrary, if

Figure 1. Above median whites and right tail blacks.



anything they appear somewhat more prone to transition to marriage than whites. We might ask why the race effect is reversed, albeit to levels that are not strongly significant. It seems reasonable to speculate that this reflects the Wilson and Neckerman (1986) hypothesis. Black men who qualify in our methodology for the stronger half of the white distribution are particularly attractive in the black population. Given that the great majority of marriages are homogamous with respect to race, the relative scarcity of those men marks them strongly for marriage.

5.4. Above median blacks and left tail whites

One way to subject this idea to an indirect test is to examine two scenarios that juxtapose relatively strong black men with relatively weaker subsets of white men. The first of these is shown in Figure 2. This modifies the experiment in Figure 1 and Table 5 by combining black men located above the median of the black population and white men whose latent earnings would fall below the median if they were placed in the black population. The two groups are statistically contiguous, but favoring black men by construction. The sample sizes are 3,812 for all usual workers and 3,488 when zero earners are excluded. If the latent earnings hypothesis is plausible, estimates of the probit model should reflect a race gap that is positive for black men to an extent that exceeds what is seen in Table 5.

Results of estimation are presented in Table 6. The estimated race coefficients conform to expectations. With the latent earnings advantage contrived in favor of black men, they display a commensurately stronger propensity to marry. The estimated coefficient for all workers is 0.21 and strongly significant ($t = 2.82$). Excluding zero earners produces an estimate of 0.23 ($t = 3.07$). Both estimates are substantially larger in magnitude and more precise than their counterparts in Table 5. Estimated coefficients of the control variables are generally consistent with those in Table 5. The only exception is the homeownership variable, which remains negative but increases in magnitude and becomes significant.

5.5. Right tail blacks and left tail whites

An extension of this experiment removes the contiguous nature of the statistical marriage market and increases the disparity between high latent earning black men and low latent earning white men. This scenario is illustrated in Figure 3. Black men who qualified for the first experiment, as seen in Figure 1, i.e., strong enough to exceed the white median, form the first component of the sample. The remaining component consists of white men who qualify for the second experiment, as seen in Figure 2, i.e., located below the black median. This produces a sample of 3,731 usual

Figure 2. Above median blacks and left tail whites.

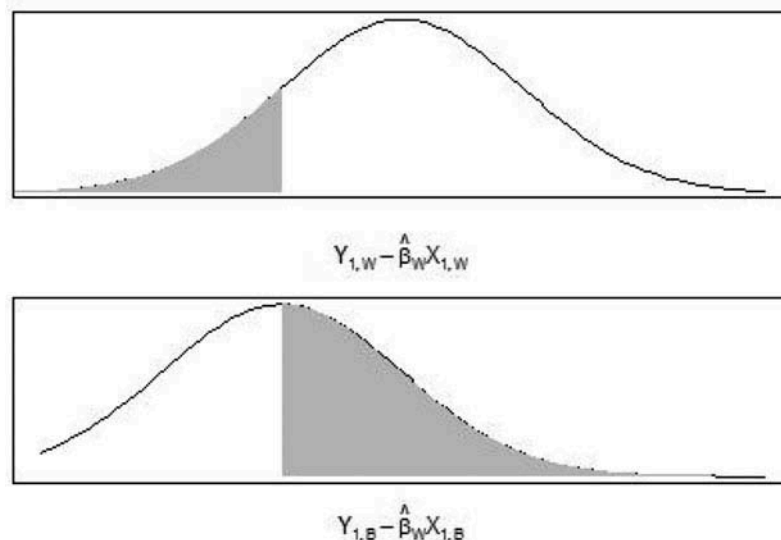
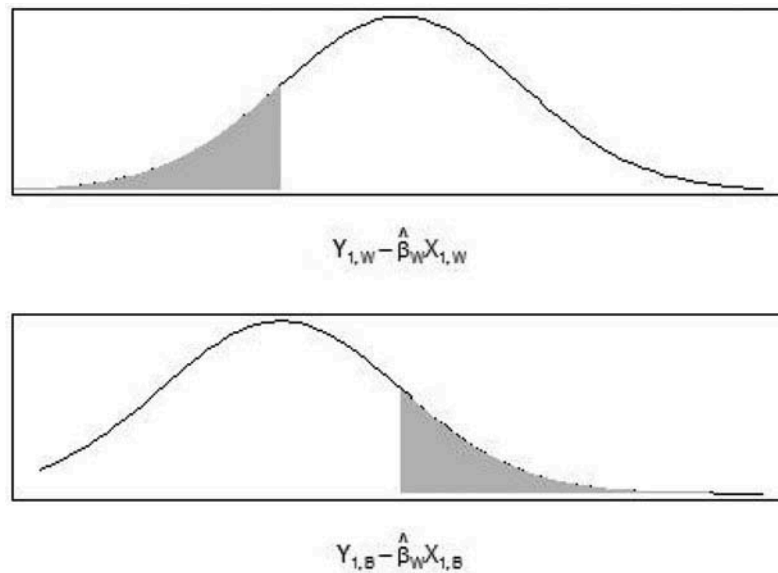


Table 6. Probit estimates: above median blacks with left tail whites

Variable	All self-reported workers	Excluding zero earners
BLACK	0.206 (2.82)	0.228 (3.07)
OWN HOME	−0.180 (2.99)	−0.184 (2.91)
AGE	0.159 (3.37)	0.142 (2.90)
AGESQ/100	− 0.003 (3.08)	−0.002 (2.59)
EDUCATION	−0.014 (1.18)	−0.016 (1.27)
DISABLED	−0.778 (2.00)	−0.498 (1.56)
YEAR1 = 2003	−0.337 (3.21)	−0.402 (3.63)
YEAR1 = 2004	−0.650 (5.32)	−0.566 (4.58)
YEAR1 = 2005	−0.576 (4.94)	−0.937 (6.32)
YEAR1 = 2006	−0.256 (4.94)	−0.217 (2.10)
YEAR1 = 2007	−0.341 (3.41)	−0.323 (3.13)
YEAR1 = 2008	−0.413 (3.99)	−0.497 (4.46)
YEAR1 = 2009	−0.477 (4.51)	−0.441 (4.85)
CONSTANT	−3.100 (4.80)	−2.637 (4.00)
N	3,812	3,488

Figures in parentheses are absolute t statistics. Dependent variable = 1 if individual transitioned to marriage, =0 otherwise.

Figure 3. Right tail blacks and left tail whites.



workers and 3,398 if zero earners are excluded. If the latent earnings hypothesis is plausible, then the estimates should reflect a black propensity that is at least as large as those in the first two experiments (Tables 5 and 6, with respective Figures 1 and 2).

The probit estimates are presented in Table 7, which confirm the intuition of the experiment. The race coefficient in column one (all workers) is 0.24 and strongly significant ($t=3.07$). Excluding zero earners reduces the estimate slightly to 0.23 while remaining significant ($t=2.91$). The relative earnings strength of these black men, combined with the relative demographic scarcity of marriageable black men in general, makes them particularly attractive as potential spouses. Estimated coefficients for the control variables are similar in magnitude and precision with their counterparts in Tables 5 and 6.

Table 7. Probit estimates: right tail blacks with left tail whites

Variable	All self-reported workers	Excluding zero earners
BLACK	0.235 (3.07)	0.229 (2.91)
OWN HOME	−0.159 (2.63)	−0.177 (2.76)
AGE	0.162 (3.39)	0.139 (2.79)
AGESQ/100	−0.003 (3.09)	−0.002 (2.49)
EDUCATION	−0.014 (1.18)	−0.014 (1.11)
DISABLED	−0.754 (1.93)	−0.485 (1.52)
YEAR1 = 2003	−0.341 (3.18)	−0.394 (3.48)
YEAR1 = 2004	−0.629 (5.11)	−0.545 (4.33)
YEAR1 = 2005	−0.551 (4.70)	−0.901 (6.03)
YEAR1 = 2006	−0.239 (2.34)	−0.188 (1.81)
YEAR1 = 2007	−0.306 (3.03)	−0.286 (2.75)
YEAR1 = 2008	−0.388 (3.72)	−0.507 (4.40)
YEAR1 = 2009	−0.454 (4.25)	−0.412 (3.71)
CONSTANT	−3.159 (5.40)	−2.868 (4.17)
N	3,731	3,398

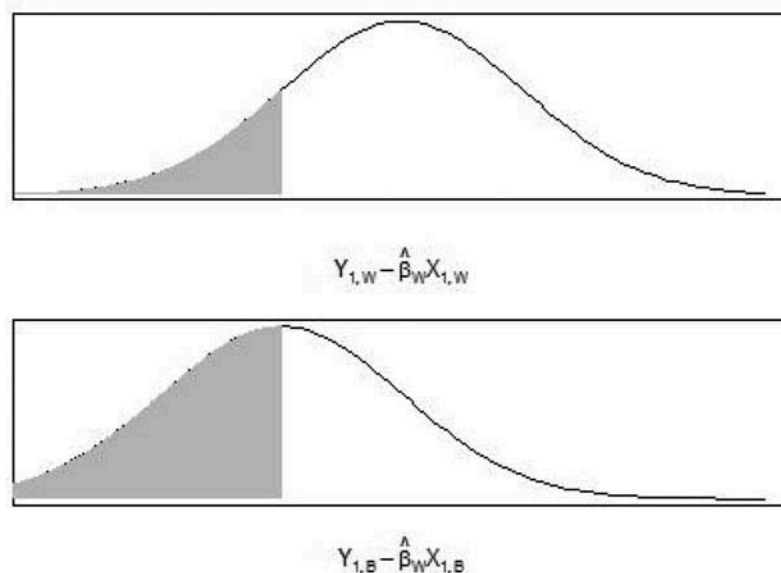
Figures in parentheses are absolute t statistics. Dependent variable = 1 if individual transitioned to marriage, =0 otherwise.

The results in Tables 6 and 7 raise a useful question for future research. In markets where the frame of reference is low earning white men, high earning black men emerge as strongly prone to marry. In these markets, with low earning whites as relatively unattractive spouses, there is a possibility that single white women might cross the race boundary into marriage with black men. Although interracial marriage is not the norm in the population of all marriages, it might occur more frequently in marriage markets with low earning white men.

5.6. Below median blacks and left tail whites

The next scenario addresses this idea in an indirect manner by combining low latent earners of both races. The sample extracts white men from Figures 2 and 3 (latent earnings below the black median) and black men from below the black median. This scenario is illustrated in Figure 4. In this

Figure 4. Below median blacks and left tail whites.



market, all men are relatively unattractive as potential spouses. The resulting samples are 3,815 for all usual workers and 3,488 when zero earners are excluded. Estimates of the probit transition model are presented in Table 8. Estimated coefficients for the control variables are again similar to those in Tables 5–7. For the first time in this series of experiments, the estimated race gap is negative. However, in both columns, it is small in magnitude and not significant (-0.01 , $t = 0.15$ in column one; -0.01 , $t = 0.08$ in column two). The evidence thus suggests that, in markets with low latent earners, transitions to marriage make little distinction based on race. This apparent absence of a race gap again confirms the latent earning hypothesis.

5.7. Entire sample: inclusion of latent earnings in the probit equation

Another way to test the latent earnings hypothesis is to include the residual measures (Equations (5) and (6) for blacks and whites, respectively) as a single race-specific

variable in the probit equation. The augmented probit equation is as follows:=

$$P(M_{i2} = 1)$$

$$P(M_{i2} = 1) = \Phi \{ \delta' z_{i1} + \lambda \bullet [BLACK \bullet (y_{i1,B} - \hat{\beta}_W x_{i1,B}) + (1 - BLACK) \bullet (y_{i1,W} - \hat{\beta}_W x_{i1,W})] \} \quad (4')$$

The motivating idea of this paper can be tested directly by using the complete sample and examining the estimate of the parameter λ in (4') in anticipation of a positive sign. Since the vector of explanatory variables includes the race dummy variable, this approach also allows us to determine whether the race gap persists across the entire sample after controlling for latent earnings.

Estimates for this experiment are presented in Table 9. Excluding zero earners, the coefficient for the composite earnings residual is positive and strongly significant. In its presence, the race gap largely disappears; although black men have a modest positive propensity for marriage, the estimate is not strongly significant. Thus, consistent with other evidence in this paper, the marriage gap appears to be an earnings phenomenon rather than something that can be strictly or directly attributed to race.

Table 8. Probit estimates: below median blacks with left tail whites

Variable	All self-reported workers	Excluding zero earners
BLACK	-0.014 (0.18)	-0.008 (0.10)
OWN HOME	-0.146 (2.38)	-0.150 (2.29)
AGE	0.195 (3.80)	0.169 (3.28)
AGESQ/100	-0.003 (3.59)	-0.003 (2.99)
EDUCATION	-0.002 (0.15)	0.008 (0.56)
DISABLED	-0.747 (1.94)	-0.444 (1.41)
YEAR1 = 2003	-0.317 (3.00)	-0.321 (2.91)
YEAR1 = 2004	-0.630 (4.91)	-0.566 (4.41)
YEAR1 = 2005	-0.486 (4.15)	-0.796 (5.49)
YEAR1 = 2006	-0.205 (1.99)	-0.123 (1.19)
YEAR1 = 2007	-0.282 (2.74)	-0.265 (2.47)
YEAR1 = 2008	-0.335 (3.18)	-0.487 (4.15)
YEAR1 = 2009	-0.381 (3.58)	-0.389 (3.51)
CONSTANT	-3.729 (5.40)	-3.588 (5.06)
N	3,815	3,488

Figures in parentheses are absolute t statistics. Dependent variable = 1 if individual transitioned to marriage, =0 otherwise.

Table 9. Probit estimates: entire sample with earnings residual in the model

Variable	All self-reported workers	Excluding zero earners
BLACK	0.039 (0.76)	0.059 (1.10)
LATENT EARNINGS: $BLACK \bullet (y_{i1,B} - \hat{\theta}_W x_{i1,B}) +$ $(1 - BLACK) \bullet (y_{i1,W} - \hat{\theta}_W x_{i1,W})$	0.008 (1.02)	0.100 (4.10)
OWN HOME	-0.085 (2.13)	-0.081 (1.96)
AGE	0.145 (4.79)	0.159 (5.00)
AGESQ/100	-0.002 (4.35)	-0.003 (4.57)
EDUCATION	-0.009 (1.30)	-0.007 (0.87)
DISABLED	-0.659 (3.02)	-0.626 (2.84)
YEAR1 = 2003	-0.401 (5.78)	-0.391 (5.48)
YEAR1 = 2004	-0.478 (6.34)	-0.464 (6.00)
YEAR1 = 2005	-0.499 (6.48)	-0.572 (6.91)
YEAR1 = 2006	-0.258 (3.80)	-0.249 (3.55)
YEAR1 = 2007	-0.317 (4.73)	-0.291 (4.22)
YEAR1 = 2008	-0.376 (5.37)	-0.386 (5.31)
YEAR1 = 2009	-0.458 (6.37)	-0.462 (6.20)
CONSTANT	-2.998 (7.18)	-3.236 (7.37)
N	8,381	7,849

Figures in parentheses are absolute t statistics. Dependent variable = 1 if individual transitioned to marriage, =0 otherwise.

6. Conclusion

There is some intuitive appeal in the idea that the marriage race gap arises from a deficiency in the supply of marriageable black men relative to black women to an extent not found in the white population. Yet the literature has not produced consistent evidence to that effect. The conventional approach has been to view marriage markets as local or regional phenomena. Using a variety of metrics and index measures that act as controls for regional supplies of potential spouses, a common conclusion has been that the race gap persists even in the presence of those controls. Examples include Lloyd and South (1996, p. 1174), Lichter et al. (1992, p. 796), Raley (1996, p. 1115), Wood (1994, p. 188), and Brien (1996, p. 773).

These findings have led some researchers to speculate that the marriage gap arises in part due to fundamental racial disparities in the culture of marriage and propensities to enter marriage. The following statements are representative of that view:

Schoen and Kluegel (1988, p. 904): ... improved economic conditions, combined with negative attitudes toward marriage formed during childhood poverty, and dissonant beliefs about male and female roles may have led to greater decline in black marriage propensities.

Lloyd and South (1996, pp. 1115, 1116, emphasis added): ... the racial difference in men's marriage propensities admits no simple explanation ... we suggest that deteriorating economic prospects generate a rejection of marriage and that, because of strong peer group ties these variables *diffuse throughout the population of young black men*.

Seitz (2009, pp. 410, 411): The results ... suggest that black women receive significantly less utility from marriage than white women.

The state of these findings has led some researchers to advocate alternative empirical approaches. Brien (1996, p. 773), for example, suggests that the scope of an individual's marriage market might be the result of endogenous rational choice, with the implication that future research should reconsider what it is that constitutes an individual's marriage market. It is this idea that motivates

the present study, along with the suggestion by Keane and Wolpin (2010, p. 863) that black women face different distributions of potential husbands' earnings than do white women. Their counterfactual experiments that equalize those distributions tend to produce convergence in white-black marriage rates (Keane & Wolpin, 2010, pp. 867, 868).

Accordingly, this study approaches the marriage market from a different perspective, moving from geographic location to statistical position in the distribution of earnings. In particular, we focus on unmeasured or latent aspects of earnings. Our empirical approach is to estimate earnings equations for young single white men. Imputing those estimated parameters into the earnings distribution for unmarried black men, we obtain regression-adjusted earnings residuals as individual-specific indexes of latent earning capacity. Equalizing the earnings parameters allows us to counterfactually place each black man in the white distribution of latent earnings, and we define the marriage market in terms of that location. Essentially, spouses are matched not geographically but rather within tranches of the latent earnings distribution.

A key point, and we believe one with intuitive appeal, is that traits which are ostensibly latent to the econometrician are observable or at least subject to assessment by potential spouses. These include ambitions and aspirations, general demeanor and character, and even physical appearance, which has been shown to correlate with earnings (Hamermesh & Biddle, 1994). Consequently, latent attributes that affect earnings might also induce transitions to marriage.

Our results are informative. In a series of experiments that juxtapose selected subsets of black and white men at several points of interest in the earnings distribution, we find that the regression corrected (via probit models) race gap is altered substantially. For example (Table 5 and Figure 1), if we place white men exceeding the white median with black men who likewise exceed the white median, holding constant across races the skill prices associated with earnings, transition to marriage slightly favors those black men. In other contrived samples, the race gap is further reversed from where it has been reported in most of the literature. For example (Table 7 and Figure 3), black men located above the white median are strongly more prone to marry than white men located below the black median.

Another approach, which differs from experimentally paired samples, is to simply include the latent earnings index directly in a model of transition to marriage for the entire sample. In that case (Table 9), strong latent earnings appear to significantly induce marriage. More importantly, the race gap becomes insignificant in that model.

The implication of these findings is that the race gap might not be due to race *per se*, nor to cultural norms that differ between races, but rather to fundamental differences in latent earning capacity. A useful research question for future work is the extent to which single women are discerning of those latent skills. Intuition suggests that they are at least somewhat adept at those assessments. If we accept the premise that discernment of quality mates is more efficacious among "high quality" women, as defined by schooling or earnings, then we should observe a higher propensity to marry among high-quality black women relative to the general black population. There is some evidence of a suggestive and indirect nature to that effect (Bennet et al., 1989; Lichter et al., 1992; Keane & Wolpin, 2010). Moreover, marriage propensities of poor black women do not appear to differ significantly from those of poor white women (McLaughlin & Lichter, 1997). Given that marital matching continues to occur predominantly within race, and together with the findings of this paper, this suggests that there is relatively strong matching on the basis of earnings quality in the black population. The importance of this issue relates to the distribution of income and earnings within the black population. If high earners tend to match with high earners, then over time the process of spouse selection will tend to bifurcate the black population at the extremes of the earnings distribution.

To the extent that marriages sometimes match individuals of different races, the results of this study suggest that the occurrence of race mismatches might not occur uniformly across the

earnings distribution. Evidence by Gullickson (2006) is suggestive in that regard. He reports that black man and black women with low levels of education are less likely to enter interracial marriages than more highly educated blacks. The implication is that low-educated blacks are constrained in the interracial marriage market.

It is worthwhile to recall that the Wilson and Neckerman (1986) hypothesis has earned support, albeit limited and cautionary, in previous research. Consequently, another useful endeavor for further research might be to combine a distribution-based approach like that used in this study with the region-oriented method that has characterized the literature.

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Correction

This article has been republished with minor changes. These changes do not impact the academic content of the article.

Notes

1. *Annual Social and Economic Supplement of the Current Population Survey*, Table MS-1. <http://www.census.gov/hhes/families/data/marital.html>.
2. The survey in this section is devoted to studies that provide primary motivation for this paper. Additional surveys are found in Wood (1995) and Brien (1996).
3. As described in Section 4, the sample is limited to black and white men, all of whom are never married in year one.
4. This approach of imputing white coefficients in an earnings equation for black wage earners has been used by Darity et al. (1998), who study race-based earnings disparities as a function of household family structures.
5. This is borne out by the descriptive statistics for the earnings residuals. For the entire sample, the average residual for white men, calculated using Equation (6) is, as expected, very close to zero (-8.24×10^{-10}). By contrast, the average residual for black men, calculated from Equation (5), is -0.101 .
6. It is worth emphasizing that previous studies have found significant race gaps even after controlling for extensive sets of background variables. See, for example, Lichter, McLaughlin, Kephart, and Landry (1992), Brien (1997), and Lloyd and South (1996).

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