Sex Ratios and Risky Sexual Behavior

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Abstract

Black Americans have dramatically higher rates of sexually transmitted infections (STI), including HIV/AIDS, despite constituting a mere 12 percent of the population. Epidemiologists have suggested that these racial disparities persist because of a greater degree of concurrency in Black sexual networks, but this invites a question: why is the degree of concurrency higher in Black sexual networks? In this paper, I emphasize the relative shortage of men in Black communities, created largely by the high rates of Black male incarceration. I hypothesize that these high "sex ratios" allow for men with tastes for sexual diversity to form concurrent partnerships, as well as affects their condom use.

I test this hypothesis using data from the 2000 Census and the National Longitudinal Survey of Youth (1997). My identification strategy exploits the fact that the overwhelming majority of sexual relationships occur between men and women of similar age, race and geographic location. I first examine the effect of the sex ratio on concurrency by focusing on its effect on men at various points of the sex partnership distribution. Next, I investigate the sex ratio's effect on condom use. A surplus of women both improves a man's ability to negotiate sex without condoms, but if it increases the degree of concurrency in the network then it also increases the risks associated with unprotected sex. Thus, the effect of the sex ratio on condom use is theoretically ambiguous.

I find that a 1-point increase in the ratio of women to men ($\times 100$) causes Black males at the .90 quantile to have between .096 and .166 more partners a year. Given a 28-point change in the sex ratio, back-of-the-envelope calculations put this at between 2.7—4.6 additional partners a year. Furthermore, I find that men at the .10 quantile of the condom use distribution reduced condom use 1.5 points for every 1 point change in the ratio of women to men, while men at the median increased their condom use 1 point.

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1 Introduction

Acquired immune deficiency syndrome (AIDS) has grown into a global pandemic. An estimated 38.6 million people are currently living with HIV/AIDS and more than 25 million have died from the infection since its official recognition in 1981. Although most of these infections are in Africa, as of 2003, nearly one million people had been infected by HIV/AIDS in the United States alone, and over 500,000 had died (UN 2006).

Behind the aggregate statistics is a picture of remarkable racial disparity in HIV/AIDS, and sexually transmitted infections (STI) in general. According to 2005 Vital Statistics, HIV was the seventh leading cause of death among Blacks but only the twenty-second leading cause of death among Whites. Focusing more narrowly on where the incidence falls, AIDS was listed as the leading cause of death among Black women aged 25-34 years and ranks in the top 3 causes of death for Black men aged 25-54.

According to the 2000 Census, Black Americans constituted 12.3 percent of the US population. However, Black Americans accounted for 50 percent of estimated new HIV/AIDS diagnoses. Table 1 presents AIDS case-rate data from the CDC HIV Surveillance Report for 2003. Black infections outnumbered White infections 21,174 to 12,175, with a Black case rate (per 100,000) of 75.2 and a White case-rate of 7.2. There were 10,450 infected White males, 1,725 infected White females, 13,624 infected Black males and 7,551 infected Black females. Furthermore, the case rates were 12.8 (per 100,000) for White males, 2.0 for White females, 103.8 for Black males and 50.2 for Black females.

Blacks account both for more new cases and overwhelmingly more of those arising through heterosexual contact. Table 2 reports the numbers of adults and adolescents living with HIV/AIDS at the end of 2003, by race, sex and exposure category. Eighty-five percent of all infected White males were exposed to HIV/AIDS either through exclusive male-to-male sexual contact or a combination of male-to-male sexual contact and injection drug use. According to the CDC, only 5 percent of infected White males contracted the infection through heterosexual contact. Black males, on the other hand, report much higher rates of heterosexual transmission with 22 percent claiming that as the point of exposure. Similar exposure disparities exist for Black and White women. In 2003, 64 percent of White women with HIV/AIDS were infected through heterosexual contact, compared with 75 percent Black women, and the case rate for Black women is 25 percent higher.

Epidemiologists do not fully understand the causes of these racial disparities (Adimora and Schoenbach (2005)), but several explanations have been proposed. Recent research has suggested the racial disparity in HIV/AIDS is the result of compositional differences in sexual networks (Laumann and Youm(1999)). For instance, individuals with relatively high numbers of sexual partners, or what is also called the "core group," in White sexual networks mix less often with the rest of the population, whereas the core group in Black sexual networks mixes more regularly with the fringe nodes. To the extent core-fringe mixing occurs, disease will spread more broadly through the population. Adimora et al. (2001, 2004) have emphasized the number of concurrent relationships in a sexual network as a primary factor. Concurrency, the temporal overlap of sexual partnerships, facilitates STI growth and transmission throughout a population by decreasing the average distance between infected and non-infected individuals, even holding constant the average number of sexual partnerships (Morris and Kretzschmar (1995, 1997)). Thus, small increases in the share of concurrent relationships can have significant multiplier effects on the infection's growth rate. They show that when one-half of all partnerships in a population are concurrent, the size of the epidemic after 5 years is 10 times larger than if all partnerships were sequentially monogamous.

Concurrency is hard to measure, but I am aware of at least two studies that document racial differences in concurrency patterns. Adimora et al. (2004) find higher rates of concurrent sexual partnerships among heterosexual African-Americans, aged 18-59, in rural North Carolina (a region with unusually high STI rates). Using the 1995 wave of the National Survey of Family Growth, Adimora et al. (2002) report that Black females are more likely to have had a concurrent sexual partner over the last year compared to White and Hispanic females.

Citing concurrency as a proximate cause begs a question: why do Blacks have more concurrent sexual encounters? In this paper, I emphasize the relative shortage of men in Black communities. The sex ratio, as quantified by the relative number of non-institutionalized females to males for a race-, age-, and geographic-specific "relationship" market, tends to be higher for Blacks because of high levels of Black male incarceration (Raphael (2004) and Charles and Luoh (2006)). I am not the first to propose a connection between the sex ratio and concurrency. For example, Adimora et al. (2002) and Adimora and Schoenbach (2005) argue that a shortage of men may cause women to feel desperate about the prospects of finding a stable partner, encouraging short-term relationships with less-committed mates, but their empirical evidence is largely qualitative. I am, though, the first to test for such a relationship econometrically using nationally representative data.

Closely related to my work is a study by Johnson and Raphael (2006), who attempt to directly link incarceration and AIDS outcomes among Blacks. Using aggregate data from the 1970-2000 Census and AIDS case data from the CDC, the authors present evidence that incarceration dynamics account for the racial difference in AIDS outcomes among females. Their findings suggest two separate mechanisms for disease transmission. First, by removing large numbers of males from the population, incarceration hampers the bargaining position of females in the relationship market. Second and separately, incarceration exposes males to STI risk via incapacitation. Prisons and jails have both higher STI rates and increased risky sexual behavior between male inmates. Upon release into the community, former inmates place females at risk as they re-enter the relationship market. However, because HIV incubation spells are so long (the median spell is 9 years), Johnson and Raphael (2006) are unable to distinguish between these two mechanisms.

My focus on risky sexual behavior allows me to explore the first mechanism. Utilizing data from the 1997 National Longitudinal Survey of Youth (NLSY97), I first examine the effect of the sex ratio on concurrency. Next, I investigate the sex ratio's effect on defensive STI choices, as represented by condom use. The direction of the relationship is theoretically ambiguous. On the one hand, a shortage of males improves a man's ability to negotiate sex without condoms. On the other hand, sexual behavior can have negative externalities throughout the network if it increases the probability of others matching with an infected agent (Jackson (2005) and Ballester et al. (2006)). A rational male will consider disease prevalence in deciding whether to use a condom. Therefore I estimate the effect of sex ratios on condom use separately for those with lower and higher risks of infection. Broadly, I find strong evidence that the sex ratio increases recent partners, and by extension, concurrency. The evidence is somewhat weaker that sex ratios affect condom use. Taken together, my findings provide additional support for the proposition that incarceration, operating through the sex ratio, likely increases the spread of STIs in general and HIV/AIDS in particular.

The paper is organized as follows. Section 2 reviews the racial differences in sex ratios and section 3 explores the relationship between STI prevalence and optimal STI risk. Section 4 describes my data and measures for the sex ratio, concurrency, condom usage, and the control variables, and the results from regressions; section 5 concludes the paper.

2 Racial Differences in Sex Ratios

Using data from the 2000 Census longform, I calculated White and Black sex ratios for six age cohorts: newborns, 1-5 year-olds, 6-10 year-olds, 11-15 year-olds, 16-20 year-olds, 21-25 year-olds, and 26-30 year-olds[1]. Traditionally, the sex ratio (sr) is defined as the ratio of men to women. Figure 1 depicts the *inverse*, which I utilize in estimation to facilitate interpretation of the sex ratio's effect. Figure 1 compares the (inverse) national sex ratio (i.e., sr^{-1}) for Blacks and Whites across age cohorts. The data show little racial disparity from birth or early adolescence, but by early adulthood the Black and White ratios begin to diverge sharply. This pattern is consistent with the racial differences in male incarceration reported in other studies (Raphael (2004) and Charles and Luoh (2006)). By the time Blacks reach their late twenties, the Black sex ratio has settled at .78 men for women, which translates into 128 Black women for every 100 Black men. I use a change of 28 to interpret my estimate of the partial effect of incarceration on male sexual partnerships.

The 2000 Census longform was also used to compute the ratio of non-incarcerated 18-24 year-old males to 18-24 year-old females for Blacks and Whites in all 50 states. Figures 2 and 3 show the geographic variation in Black and White sex ratios. Not surprisingly, areas where Blacks have the highest concentrations are areas with the darkest shading, representing sr^{-1} between 100 and 130. White sex ratios, on the other hand, have little to no variation in color, which is consistent with the empirical fact that White sex ratios do not vary much by state or over time.

Over 90 percent of the US Black population lives in 24 states, which have relatively high sex ratios (see Figure 4). The maps and the graphic together show a consistent pattern: by their mid-20s, Black women in every state where the Black population is relatively high (mostly the southeast) outnumber Black men in the general population

[1] The details of this calculation are discussed in section 4 below.

significantly.

3 Relationship Markets, Concurrency and Male STI Risk

Standard matching models predict that the removal of men from the marriage market reduces female probability of marriage, and empirical studies have borne this out (Brien (1997)). But there are other, more subtle, predictions that flow from such models. First, as shown by Roth and Sotomayer (1990), removing men from (or equivalently adding women to) will affect the optimal sorting of both males and females into matches. Assuming that males and females have complete ordered preferences, removing either males or females creates a new equilibrium where the gender in short supply has moved up his preference ordering, and the gender in excess supply has moved down. As a result, females will have a higher chance of matching with a male they preferred less, if they match at all. Second, females who do match are hurt because the division of output will be renegotiated in response to the removal of men (Becker (1973)). This effect induces women to make transfers to their partners because their bargaining position is weakened. The types of transfers that might take place range from the trivial, such as who picks which restaurant when going on a date, to the much more serious, such as whether the wife will tolerate sexual infidelity, abuse and other kinds of damaging behaviors. Finding a shortage of men increases concurrency can be interpreted as evidence that either the stable equilibrium has changed such that men and women with preferences for semi-polygamous matches are together and/or such "serious" transfers occur from females to males. Determining the appropriate policy response would depend on which of these is behind the equilibrium condition.

There is a large empirical literature on how marriage market environments affect relationship outcomes.[3] However, no study has examined the effect of sex ratios on sexual choices relevant to an STI epidemic, though such relationships have been suggested by Charles and Luoh (2006), Adimora, et al. (2005), and Johnson and Raphael (2006).

Sex-ratio induced concurrency raises a question about male STI risk. What effect

[3]Research has linked imbalanced sex ratios to out-of-wedlock childbirth (Wilson (1987), Willis (1999) and Neal (2004)), female-headed households (Fossett and Kiecolt (1993) and South and Lloyd (1992)), marital delays (Lichter et al. (1992) and Brien (1997)), wages and human capital investment (Angrist (2002) and Charles and Luoh (2006)), and welfare dependency (Darity (1984)).

might the removal of men from relationship markets have on contraceptive decisions. On the one hand, if some men prefer vaginal intercourse without a condom—because the sensual pleasure is greater presumably—men in a strong bargaining position might successfully negotiate less frequent condom usage, ceteris paribus. Babcock and Laschever (2003) discuss research that found women are reluctant to negotiate condom use with their husbands, since doing so was tantamount to accusing them of sexual infidelity. On the other hand, if the sex ratio increases the degree of concurrency across the sexual network, then the risk of contracting an STI has increased for all people attached to that network. The more partners a person has, the more likely they will eventually come into contact with an infected female, and thus contract an STI themselves (Holmes et al. (1999) and Finer et al. (1999)). Thus, a lower sex ratio might induce some men to wear condoms more frequently to compensate for the increased risk.

I can illustrate this possibility formally in a simple model. Assume males know the distribution of infected and non-infected females in the population and the probability, π, of matching with an infected female,

Also assume that infection leads to certain death. Then the expected utility from having sex with a randomly selected female is

$$EU = \pi(V - \delta D) + (1 - \pi)V, \tag{1}$$

where V is the value of the sexual encounter, D is the cost of death, and δ is a discount rate. An increase in the infection rate lowers the male's expected utility but whether he will have sex depends on the relative magnitudes of his valuation of sex, his life (which could be affected by labor market conditions, wages, and human capital investment), his discount rate and the probability of infection.

Now consider a costly technology that can reduce the risk of infection, represented by the cost function $C(P)$ (where $C'(P) > 0$ and $C''(P) > 0$). This technology can reduce the likelihood of infection, but is costly to use, only partially successful and has rising marginal costs. As a result, the male's problem is affected by the failure rate of the technology, $1 - \lambda$, where $0 < \lambda < 1$. His expected utility now becomes

$$EU = P[\pi(V - \delta(1 - \lambda)D) + (1 - \pi)V] + (1 - P)[\pi(V - \delta D) + (1 - \pi)V] - C(P) \tag{2}$$

The first-order condition for the male's optimization problem is:

$$C'(P) = \pi\delta\lambda D \tag{3}$$

and its interpretation is straightforward. The male chooses a level of protection, P^*, that equates the marginal costs of his contraceptive use ($C'(P^*)$) with the marginal gain in STI avoidance ($\pi\delta r D$). From (3), one can see that an increase in the proportion of infected females in the population will cause the male to lower his STI risk by selecting a higher P^*[4]. If a shortage of males increases concurrency throughout the sexual network, then the model predicts men will respond by wearing condoms more often. Since I cannot identify which men, from my sample, have tastes for unprotected sex and which men would be responsive to the increased risk of the network, I use quantile regressions to separately estimate the effect at the .10 and .50 quantile. These two quantiles are important because they represent the two extremes. The .10 quantile describes those men who never or rarely wear condoms, which I use to identify those men with tastes for unprotected sex. And the .50 quantile describes those men who nearly always wear condoms, since the .50 quantile is equal to 100 percent.

4 Data

My data come from two sources: the 2000 Census longform survey, also known as the 5-percent sample, and the NLSY97 Geocode. The Census data are used to construct the sex ratios and the NLSY97 provides information on the sexual behavior of adolescents and young adults.

4.1 Sex Ratio Construction

Although the sex ratio is easy to understand conceptually, there is no consensus on the correct way to measure it empirically. One issue concerns the relevant age cohort. Angrist (2002) staggers the male-to-female ages in such a way that men are assumed to search

[4]Equation (3) also highlights other effects worth briefly noting. First, his discount rate influences the male's contraceptive behavior. Unsurprisingly, the more the male values his future, the more cautious he becomes. Somewhat surprisingly, though, is the effect that technological improvements has on optimal STI protection. Technological advances in STI protection causes the male to increase his optimal level of STI risk, all things equal. This result, restated, says that males can have the same level of expected utility by using contraception less frequently. Since contraception is costly, he responds to the increased efficacy of technology by depending on it less often.

among women roughly two years younger than themselves. Numerous demographers and sociologists have followed this strategy (Gutenberg and Secord (1983)). But it is unclear whether even this added level of realism is correct since there is substantial dispersion around the mean age difference between partners. The relevant geographic area or market is another issue. Fossett and Kiecolt (1991) recommend constructing sex ratios defined by race, age and either MSA or county-level. However, Brien (1997) presents evidence that county- and MSA-level sex ratios constructed from Census longform data contain serious measurement error. Using the 5-percent sample at the county/MSA level creates measurement error because some cells have so few observations. The problem is exacerbated in counties/MSAs with small Black population shares. The correlation of the measurement error with race makes an analysis based on counties or MSAs unappealing.

Measurement error is not the only reason to be skeptical about a county or MSA-level analysis. One obvious adjustment to a sex ratio imbalance in your relationship market is to move your search to another neighboring market. Females who face a deficit of men in their immediate vicinity may look outside for a mate. This is much more likely to be a problem across counties and MSAs than across states. As an example, Adimora et al. (2004) report that women in North Carolina traveled to the military base, Fort Bragg, because men were lacking in their immediate communities.

So, I base my empirical work on sex ratios calculated at the state level. For any geographic area, distinguishing the non-institutional population from the population at large is key. The Public Use Microdata Samples (PUMS) for each census includes an indicator for the institutionalized and non-institutionalized population by age, race, sex and state. I construct a race-, age- and state-specific sex ratio based on the number of non-institutionalized men and women reported in the 2000 Census longform survey. As Charles and Luoh (2006) note, institutionalized mainly consists of incarcerated individuals, particularly for a cohort as young as the one I am examining.

Following Charles and Luoh (2006), I exploit the fact that the overwhelming majority of sexual relationships occur between women and men of similar age, race and geographic location. I measure the contemporary non-institutionalized sex ratio for each demographic group defined by state of residence, age group, and race/ethnic group according to the following formula:

$$sr_{a,r,j} = \frac{\sum_{a-b}^{a+b} M_{a,r,j}}{\sum_{a-b}^{a+b} F_{a,r,j}}, \tag{4}$$

where $M_{a,r,j}$ ($F_{a,r,j}$) denotes the number of non-institutionalized men (women) of age a, race r, living in state j. Using a strategy similar to Helmchen (2005), I estimate sex ratios that vary by age according to the age ranges depicted in Table 3. I will be exploiting variation in the state-age-race-cohort cell to identify the effects of the sex ratio.

4.2 Concurrency, Condom Use and Control Variables

My measures of concurrency and condom use come from the NLSY97, a household survey representative of people living in the United States in 1997 who were born during the years 1980 through 1984. I obtained the NLSY97 Geocode through special request, which includes geographic indicators for each respondent that allow me to match each respondent to a specific state.

Concurrency reflects temporal overlap in sexual partnerships. For instance, if a male had sex with Female A on June 2nd, Female B on June 5th, and Female A again on June 10th, then one might describe this as a concurrent match. Thus, the ideal data source would contain information on both the number of recent encounters and the dates of each encounter. The National Survey of Family Growth (NSFG) provides such information but it is a repeated cross-section (not a panel) and males were not surveyed until 2002. The NLSY97 is currently comprised of six waves, following the same individuals in a panel that extends through 2002. Although it does not provide information on encounter dates, the NLSY97 does report the number of recent vaginal intercourse encounters, defined in terms of the last twelve months for 1997 and since the date of the last interview for all subsequent waves. I use the recent partner data to compute annual figures for each wave.

Male NLSY97 respondents are not only asked their number of sexual encounters during the past year, they are also queried about the number of times they wore a condom and the number of times they had vaginal intercourse since the date of their last interview. As before, I construct measures of sexual precaution using the months since the date of the last interview and respondents' answers to how much sex and how many condoms they have had since the date of the last interview, then multiplied that number by 12. I combine the measures based on these questions to calculate a condom-use rate, c/s (where

c is the frequency of condom use in the last twelve months, and s is the number of vaginal sex encounters), which I use as my safe-sex measure.[5]

The NLSY97 also provides information on a number of important covariates. In particular, I control for a person's age, education level (measured as years of schooling), and family structure (proxied by an indicator of whether respondent's biological parents are still married).

For the safe-sex regressions, I add an indicator for respondent reporting 4 or more partners in the previous period ($rp4$) to control for unobserved tastes for promiscuity. If the respondent reported 4 or more partners in the previous period. I also include a control for respondent's marital status, because other studies have found the sex ratio lowers marriage rates (Brien 1997) and marriage is usually associated with a dramatic decrease in condom usage.

My sample starts with a total of 8,984 NLSY97 male respondents in 1997. From those individuals I select Black and White non-Hispanics who have complete data for all six available years of the survey.[6] My initial sample includes 1669 Whites and 601 Blacks who are observed from 1997 to 2002. Individuals in my sample were not asked questions about their sexual histories until they reached the 14, but in 1997, 12–13 year-olds were in the sample. I imputed by examining the sexual histories of these individuals when they answered the sex history questions in 1998 and 1999. Most male children 12 to 14 years-old have not yet made their sexual debut, and so respondents who reportedly were virgins at 14 were necessarily virgins when they were 12–13. I assigned values of 0 to all 12–13-year-olds who were reportedly still a virgin when they were 14. Individuals who were sexually active by age 14, and who had made their debut when they were 12 or 13, could still have partnerships imputed to their earlier histories. First, if they had only one lifetime partner, and had made their debut when they were 12 or 13, I assigned a 1 to the appropriate year. Second, if they had more than one lifetime partner, and made their debut when they were 13, I could subtract the number of partners they'd had recently from their lifetime partners, and assign the appropriate value to both age 12–13. And if individuals had lost their virginity when they were 12, I distributed an equal number of

[5] In a few instances, an individual would report wearing condoms more times than he reported having sex. These observations were dropped from the sample.

[6] I ignored Hispanic and Latinos for the simple fact that Census data on Hispanic/Latino characterization is currently very poor. Further, Hispanic/Latino relationship markets are diverse, making it even difficult to match individuals in the NLSY97 sample with an appropriate Census race category.

partners to 12–13 depending on their lifetime partners in 14.

Table 4 reports average number of partners since the date of the last interview for all individuals, by age and race, in my sample. As can be seen, there is a monotonic increase in partnerships from age 14 to 19 for Black males, and from age 14 to 21 for White males. The 12–13-year-old data has a smaller standard deviation than all other years. It is not surprising that the means for these years are so relatively small since a majority of children do not make their debut until well into their teenage years.

One noticeable difference between Blacks and Whites is that Blacks have a much higher mean number of partners than Whites throughout most of the sample, and especially during the late teenage years. Unlike other studies, the NLSY97 top-codes recent sexual partnerships at 99. When top-coding is dealt with, these figures are consistent with other data sources, such as the National Survey of Family Growth 2002 (NSFG).

Other covariates were included in order to control for potentially confounding effects on the sex ratio. The NLSY97 included the age of the respondent (calculated in months), and this measure was included so as to control for natural changes in sexual behavior that varied by age. Respondents' highest grade completed as of the date of the interview were also recorded. And a measure of family intactness was created based on questions about parents' marriage. The family background variable is a dummy variable equalling 1 if the respondent's biological parents were still married at the date of the interview. The summary statistics for the additional covariates used in the rp regressions are in Table 5.

5 Estimation and Results

My general empirical strategy is to estimate models of the form

$$y_{ijt} = \beta_1 \, r_i + \beta_2 (sr)^{-1}_{ijt} + \beta_3 (sr)^{-1}_{ijt} \times r_i + \mathbf{x}\gamma_{ijt} + \epsilon_{ijt}, \tag{5}$$

where y is either rp or c/s, r is a race indicator set to one for Blacks, sr^{-1} is the sex ratio, $\mathbf{x}$ contains the covariates, and ϵ_{ijt} is an error term that, in principle, includes individual (i), state (j), and time (t) effects. As mentioned earlier, sr enters the regression in its inverse (i.e., as the ratio of females to males), multiplied by 100. For example, for early adulthood, sr^{-1} is 128 on average for Blacks, ages 18–24. Since I predict that the disappearance of males is positively associated with increased concurrency among Black males, this formulation allows us to easily interpret an increase in sex partners in terms of an increase in the surplus of females.

I estimate the standard pooled OLS (POLS) and linear fixed effects (FE) versions of (5). The POLS standard errors were corrected for heteroscedasticity, and the linear fixed effects standard errors were corrected for within-individual clustering. In addition, I estimate the partial effect of sr^{-1} at different quantiles of the recent-partner and condom-use rate distributions. The standard errors for the quantile regressions were obtained by bootstrapping.

5.1 Effects of Sex Ratios on Recent Partners

I first estimate (5) omitting the sr^{-1} terms to highlight the racial disparities in sexual partners, conditional on age, schooling, and whether one's parents are married. The results of this exercise are reported in Table 6. The first column presents POLS estimates of the mean difference, while the remaining columns show the racial disparities at the .5, .75, and .9 quantiles. Panel A gives the results for the entire US and panel B those pertaining to the 24 states that account for over 90 percent of the US Black population. All estimates are conditional on state and year fixed effects.

Reading across the first row of Table 6, the racial disparity in rp among adolescents and young adults is clear. On average, Blacks have .85 more partners than whites; that difference is small at the median but increases to almost 2 partners at the .90 quantile. This pattern is essentially the same in the states with the highest concentration of Blacks.

The mean effects of the covariates are as expected. Recent partners increase with age (to a point); decrease with schooling; and are lower among young people whose parents are still married. However, there is some heterogeneity in the estimated coefficients across quantiles, with magnitudes generally increasing as you move into the right tail of the rp distribution. Again, the results are similar, whether the analysis is conducted for all 50 states or the subset of 24 with substantial Black populations. For most estimates, dropping individuals from the lowest Black population states has a negligible effect. The largest effect is on the effect of the household variable. When using all 50 states, the mean difference on the household variable is –.46, but falls to –.31 when I drop individuals from the states which have few Blacks in the NLSY97.

Table 7 introduces the sex ratio terms and presents the results for the entire sample.

The first four columns repeat the specifications in Table 6. The last column reports the results from the specification that adds individual fixed effects. The focus is on the Black$\times sr^{-1}$ interaction and in every case it is positive, though not always significant or large in magnitude. The POLS estimate of the interaction coefficient is .005 and .029 for the sex ratio coefficient but neither are statistically significant. Qualitatively, the FE results are consistent with POLS, although the sr^{-1} coefficient is more precisely estimated.

When I examine sr^{-1}'s effect at different points in the rp distribution, the story is different. The effect of the sex ratio at each of the quantiles, on the other hand, is positive and significant for all the interactions. The effect at the median is .015 for Black males with a significance at the 1 percent level. Using a 28-point change in the sex ratio^{-1}, this translates into approximately 0.392 more partners a year for the median Black male. Some of this may be capturing a slightly earlier sexual debut for Black males, as well as shorter sexual relationships among Blacks. Moving higher up the quantiles, the effect of the sex ratio on Black males is more pronounced. We find that the effect is largest at the .90 quantile—the net effect of incarceration adds 2.74 more partners per year to men at the .90 quantile (.098$\times$28). In contrast, except for the weak evidence supplied by FE results, the sex ratio does not seem to matter at all for Whites, which makes sense because there is so little variation in the White sex ratio.

Given the quantile results, the failure to find a large interaction in the conditional mean is surprising. Geographic regions with small Black populations yield demographic cells that are especially sensitive to changes in the denominator and numerators. Measurement error could explain why the interaction is biased towards zero. Therefore, we dropped all respondents except for those living in the 24 highest Black population states. Table 8 presents the results for those states.

When I focus only on those 24 states, the magnitudes and the significance increases for all 5 models. Notably, the magnitudes at the three quantiles are larger and more precise. The coefficients for the .75 and .90 quantile are sizable—.068 for Black males at the .75 quantile and .166 at the .90 quantile. Back-of-the-envelope calculations suggest the sex ratio has a sizable impact on the far right tail of the recent partner distribution—Black males at the .75 quantile have approximately 2 more partners a year, while those at the .90 quantile have closer to 5 more a year, ceteris paribus.

While there is broad support that sex ratios increase concurrency, with mild effects at the central tendencies, measurement error may be responsible for some of these results. When the NLSY97 interviewer collects data on sexual history, the respondent is provided a self-administered questionnaire which is completed privately and away from the interviewer. This approach has been found to increase response rates and presumably yields more accurate results for sensitive areas such as sexuality and reproductive health. But it is also possible that other sorts of errors appear at this stage because of the lack of professional oversight. For instance, I found 29 sexual histories that showed dramatic spikes in the number of partners reported in a single year, suggesting a mistake was made by the respondent in one of the years. An example of this is when a respondent provided single digit (and some zeroes) answers to the number of partners he had since the date of last interview except for one year, and in that one year, the respondent reported 99 partners. I identified numerous examples where individuals reported outlier events which we have labeled "|DRP|>40". These are individuals who reported a difference in the number of partners that exceeded 40 in absolute value. After dropping these 29 respondents, I reran all the models to see whether these improbable events were driving the results. Those estimates are reported in Table 9.

Removing those individuals causes all results to become statistically significant, including the linear FE which is now significant and positive at the 5 percent level. The POLS result is smaller in magnitude, as is the effect at the .75 and .90 quantiles. Furthermore, the coefficient on the .75 quantile is no longer significant for White males.

Finally, I addressed one more potential source of bias in my estimates—the imputation of partnerships to 12–13-year-old. While I correctly imputed values for everyone I could identify, it is possible that I have over-represented virgins in my sample. If so, then this could artificially cause the sex ratio to pick up an effect that is purely spurious. To check the robustness of my results, I dropped the 1997-1998 waves and rebalanced the panel for 1999-2002. This allows for me to observe all individuals who provided sexual data and who were in the survey all four years. But, doing this also causes me to increase every individual in the sample's starting age by two years. Since sexual debut occurs on average around 15 for most males, this necessarily causes the panel to begin after most men have already made their debut, and thus does not allow for sexual debut for identification. Nevertheless, I should still find an effect at the upper quantiles, since individuals with tastes for promiscuity still should show a responsiveness to the changes in their mating options, independent of others' decisions to make their debut slightly earlier. Table 10 shows the effects of the sex ratio on individuals in the highest Black population states, with the most volatile sexual histories dropped, and for the 1999-2002 balanced panel.

Unsurprisingly, the coefficient on the sex ratio at the mean and median is no longer statistically significant. The coefficient on the interaction term at the .90 quantile, on the other hand, remains positive and statistically significant at the 10 percent level (P-value approximately .087).

Finally, I note the adding the sex ratio amplifies the estimated racial disparity in recent partners, but changes the covariates' coefficient estimates very little. However, in the presence of individual fixed effects, the covariates (other than age), explain little of the variation in recent partners. In the presence of individual fixed effects, age is negative and statistically significant.

5.2 Effects of Sex Ratios on Condom Use

In section 3, I discussed two possible behavioral responses to changes in the sex ratio. On the one hand, men with tastes for unprotected sex could negotiate less frequent condom use if their bargaining power is a function of their options on the relationship market. In which case, one might find a negative relationship between the sex ratio and their condom use. On the other hand, if the sex ratio causes the degree of concurrency to increase, then it increases STI risk within the network. This would lower expected utility and cause men to increase their condom use to compensate for the increased risk. Theoretically, therefore, the effect of the sex ratio on condom use is ambiguous.

Table 11 reports summary statistics for the covariates at the mean, the various quantiles, and the variable's standard deviation based on the balanced, all wave, highest black population state datasets. Like with the previous section, I used several different slices of the data. Since the results were the same regardless of the slice used, I will report the results for the all waves, highest black population states with the "sexually volatile" individuals removed from the sample. Since condom-use is measured as the ratio of condoms (c) to sexual encounters (s), the sample is necessarily a subset of sexually active individuals. As can be seen, over the entire period, of those males who were sexually active, respondents wore condoms during vaginal intercourse approximately 69 percent of the time.

Table 12 reports the racial differences in condom use controlling for familiar regressors such as age, education, and family background. New to these regressors are marital status and the sexual activity from the previous period. I include a control for marital status because of evidence that sex ratios cause Blacks to marry later (Brien 1997). Since marital status is also a situation wherein condom use declines considerably, it is important to control for this omitted variable. I also include controls for the previous period's sexual activity—an indicator variable equalling 1 if the person reported an annualized number of partners of 4 or more the previous year. I include this control so as to capture the person's overall level of promiscuity, which is presumably exogenous to contemporaneous condom use.

The results show that Black males have considerably higher condom-use rates compared to White males. Black males have a condom use rate that is 14.5 points higher than the corresponding White male on average. This difference varies significantly across the condom use distribution, with smaller racial differences at both the .10 and .50 quantile. In all cases, the racial difference in condom use is precisely measured. From this table, we also see that condom-use declines with age, but that again the effect is heterogeneous along the condom use distribution. Each month the male reduces condom use by half a point at the conditional mean. This effect, again, is stronger at the .25 quantile than at the .10 or .50 quantiles.

Education is also associated with an increase in condom use. For every grade completed, males reportedly increase their condom use by roughly 2.6 points on average. Except for the .10 quantile, the effect is positive and statistically significant at the 1 percent level for the other estimates as well. Family background is also associated with an increase in condom use, except at the .10 quantile. And consistent with our predictions, marital status is associated with a dramatic decrease in condom use, except for people at the .10 quantile which is small in magnitude and statistically insignificant. The inclusion of a control for the previous period's sexual activity is both small in magnitude and statistically insignificant.

Table 13 reports the results of controlling for the sex ratio in the respondent's state of residence. Inference is mainly on the interaction of race with sr^{-1}. The coefficient on the POLS and FE regressions reveal little support for my hypotheses. Of course, the mean effect could be masking differential behavioral responses across the condom-use distribution. The median quantile gives some evidence for this. The estimated coefficient of the $Black \times sr^{-1}$ interaction is approximately equal to and highly significant, suggesting that the median effect of the sex ratio on Blacks is to increase their condom use.

To test for whether men with "tastes" for unprotected sex use their advantageous position in the relationship market to negotiate less frequent condom use, I estimated the effect of the sex ratio on condom use at the .10 quantile in the condom use distribution. Table 13 reports that the effect of the sex ratio on Black male condom use at the .10 quantile was negative and significant—condom use decreased 1.5 points for every 1 point change in the sex ratio for Black males.

6 Conclusion

The goal of this paper has been to identify a causal link between the sex ratio and risky sexual behavior among men in the form of concurrency and condom use. Using data from the Census longform and the NLS97, I present strong evidence that, for Blacks, recent sexual partners (my measure of concurrency) increases with the surplus of females for Blacks. On average, moving from parity (100) in the sex ratio^{-1} to the average surplus of women among 18–24-year-old Blacks (128) translates into an increase in Black male partnerships of around 1 additional partner a year for most of my specifications. At the .90 quantile of the recent-partner distribution, the effect consistently enters positive and significant. Depending on the slice of the data used, the net effect is between 2.7-4.6 extra partners. The support for a link between the sex ratio and condom use is much weaker, but what there is suggests the possibility that Blacks adjust to the greater risk of matching with an infected female when there is a surplus of females by increasing condom use. I also find some evidence that men in the left part of the condom-use distribution respond to increasing numbers of unattached women by reducing their condom use rate. This, I argue, supports my conjecture that men with tastes for sex-without-condoms respond to the increased bargaining position by negotiating a lower condom-use rate. Overall, my findings imply that incarceration, by creating a shortage of Black males in relationship markets, likely increases concurrency and therefore the spread of STIs.

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Table 1 Estimated US AIDS cases and rates by race and gender, 2003

Ethnicity	Infections	Rate
White (non-Hispanic) Males	10,450	12.8
White (non-Hispanic) Females	1,725	2.0
White (non-Hispanic) Total	**12,175**	**7.2**
Black (non-Hispanic) Males	13,624	103.8
Black (non-Hispanic) Females	7,551	50.2
Black (non-Hispanic) Total	**21,174**	**75.2**

Source: CDC HIV Surveillance Report 2003

Table 2 Estimated numbers of adults and adolescents living with HIV/AIDS at the end of 2003 by race, sex, and exposure category

Exposure Category	Infected	Percent	Infected	Percent
Male	**White**		**Black**	
Male-to-male sexual contact	86,674	76	50,675	47
Injection drug use	10,550	9	23,658	22
Male-to-male sexual contact and inject drug use	10,431	9	7,817	7
Heterosexual contact	5,178	5	23,513	22
Other	1,524	1	1,198	1
Subtotal	**114,358**	**100**	**106,861**	**100**
Female	**White**		**Black**	
Injection drug use	6,625	34	13,244	23
Heterosexual contact	12,494	64	43,957	75
Other	447	2	1,118	2
Subtotal	**19,566**	**100**	**58,319**	**100**

Source: 2003 CDC HIV Surveillance Report

Table 3 Definition of Age-specific, Race-specific sex ratios

Male Ages	**Searching Among Women Ages**	**Competing with Men Ages**
a$\in [13, 19]$	[a-1, a+1]	[a-1, a+1]
a$\in [20, 24]$	[a-2, a+2]	[a-2, a+2]

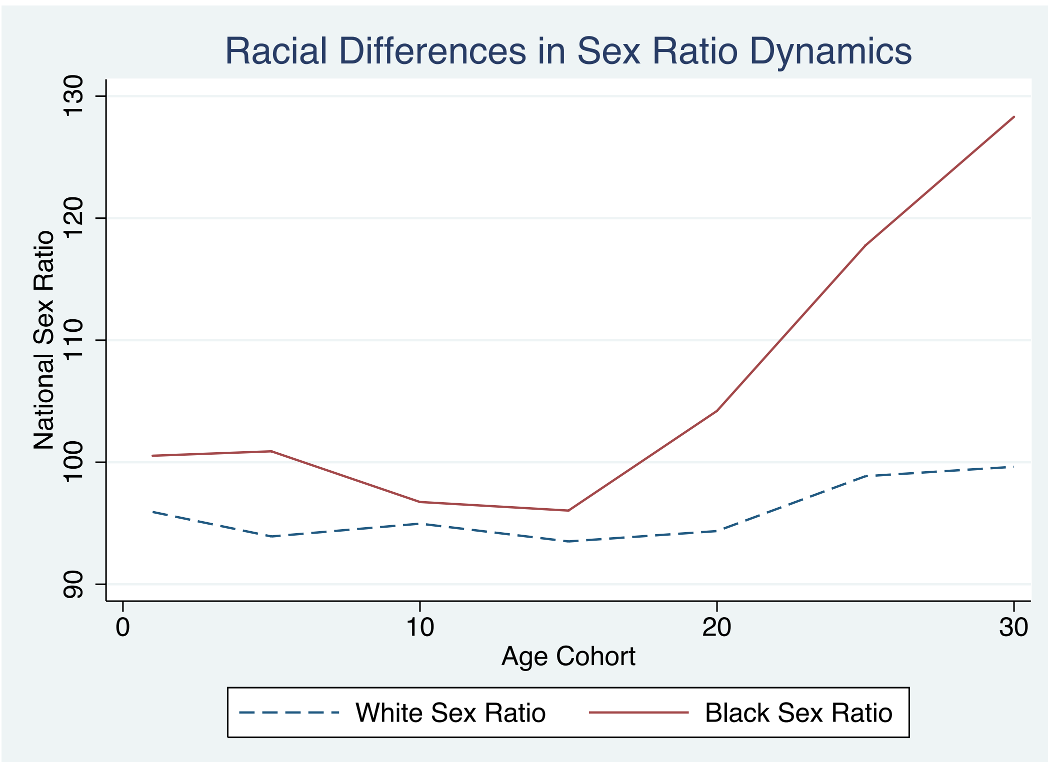

Figure 1 U.S. Sex Ratio^{-1} by Race and 5-Year Age Cohorts

Table 4 Summary Description of Recent Partners at Points in the Distribution

Black Males, Balanced Panel, All Waves, All States

	12 yo	13 yo	14 yo	15 yo	16 yo	17 yo	18 yo	19 yo	20 yo	21 yo	22 yo	23 yo
Mean	0.09	0.08	0.55	1.02	1.49	2.39	3.17	3.60	3.34	3.34	3.45	1.60
SD	0.41	0.46	2.93	3.19	3.05	6.62	7.97	9.35	6.24	6.98	6.96	2.07
Median	0	0	0	0	0	1	1	1	1	2	1	1
75th perc.	0	0	0	1	2	2	3	3	3	3	4	2
90th perc.	0	0	1	3	4	5	6	8	8	7	8	5
N	69	108	241	368	470	539	608	514	427	261	117	10

White Males, Balanced Panel, All Waves, All States

	12 yo	13 yo	14 yo	15 yo	16 yo	17 yo	18 yo	19 yo	20 yo	21 yo	22 yo	23 yo
Mean	0.05	0.00	0.12	0.14	0.85	1.21	1.59	1.74	1.84	2.15	2.13	1.43
SD	0.27	0.05	1.36	0.66	4.58	5.05	5.26	5.47	4.57	5.11	5.62	1.45
Median	0	0	0	0	0	0	1	1	1	1	1	1
75th perc.	0	0	0	0	1	1	2	2	2	2	2	2
90th perc.	0	0	0	0	2	3	3	4	4	5	4	3
N	233	363	708	1069	1351	1494	1639	1321	1019	655	292	14

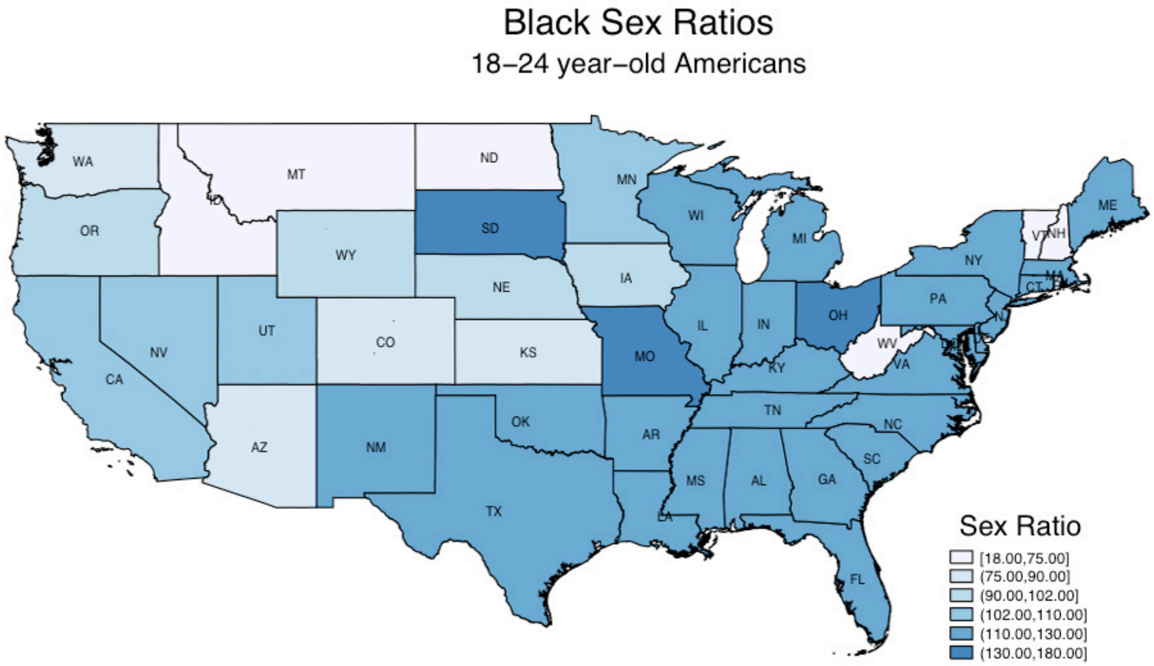

Figure 2 Black 18-24 Year olds State Sex Ratios^{-1} by Race

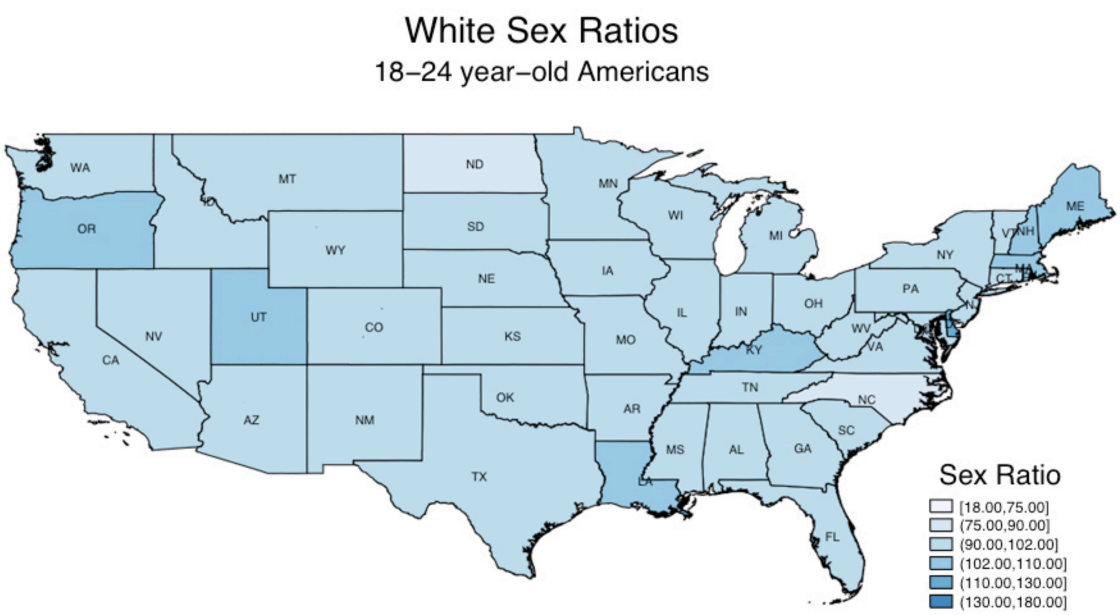

Figure 3 18-24 White Year olds State Sex Ratios^{-1} by Race

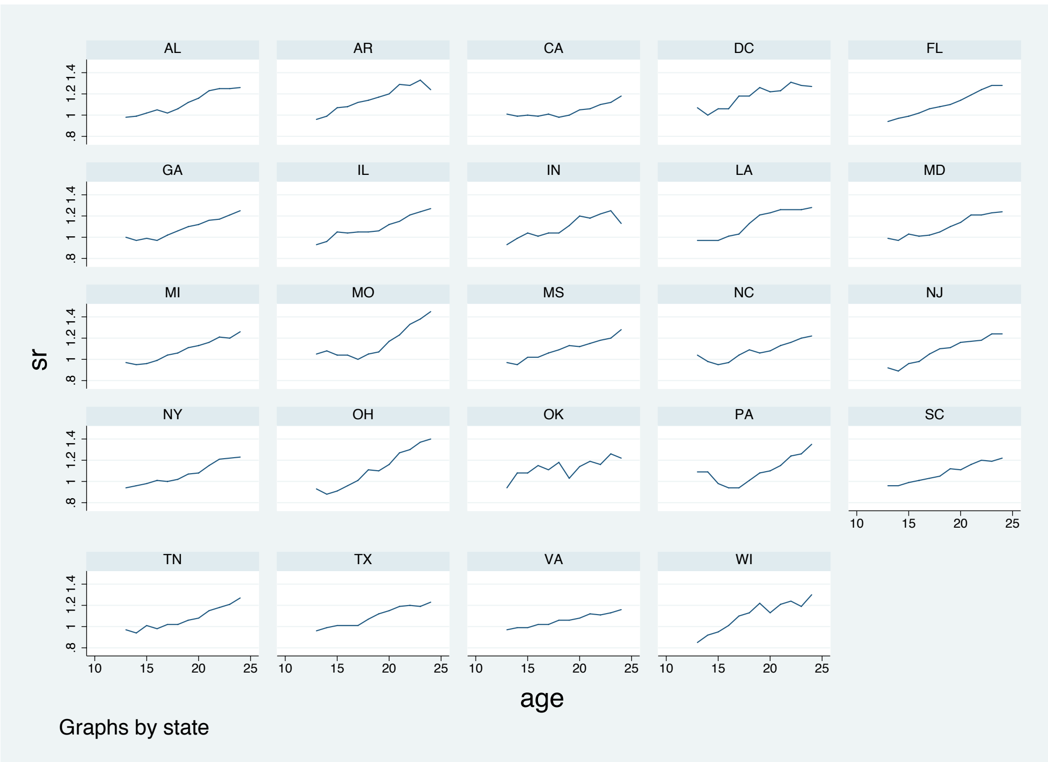

Figure 4 Black Sex Ratios for the 24 Blackest States (Source: Census 2000 Longform Survey)

Table 5 Summary Statistics for Covariates

Balanced, All Waves, All States						
	Mean	SD	.50	.75	.90	N
Age	214.330	28.263	215	235	252	13620
Highest Grade Completed	10.319	2.144	11	12	13	13620
Both Parents Married	.477	.499	0	1	1	13620

Table 6 The Effect of Race on Sex Partners (Corrected SE)

A. Balanced, All Waves, All States				
	POLS	.50	.75	.90
Black	.853	.134	.792	1.889
	(.136)	(.018)	(.089)	(.192)
Age	.045	.015	.040	.075
	(.005)	(.001)	(.002)	(.005)
Highest Grade Completed	-.282	-.069	-.218	-.500
	(.059)	(.009)	(.020)	(.050)
Both Parents Married	-.459	-.124	-.224	-.392
	(.087)	(.019)	(.037)	(.070)
Year Effects	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes
R^2	.05	.066	.101	.114
N	13542	13542	13542	13542
B. Balanced, All Waves, Highest Black Pop. States				
Black	.844	.115	.787	1.916
	(.146)	(.024)	(.067)	(.166)
Age	.049	.014	.041	.077
	(.006)	(.002)	(.003)	(.005)
Highest Grade Completed	-.311	-.066	-.232	-.508
	(.068)	(.011)	(.033)	(.053)
Both Parents Married	-.311	-.107	-.174	-.304
	(.101)	(.023)	(.024)	(.077)
Year Effects	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes
R^2	.05	.062	.098	.113
N	9956	9956	9956	9956

Table 7 Estimated Effect of Sex Ratio on Male's Recent Sex Partners (Corrected SE)

Balanced, All Waves, All States					
	POLS	.50	.75	.90	Linear FE
Black	.036	-1.391	-3.842	-8.218	(dropped)
	(2.684)	(.496)	(1.092)	(1.933)	
Sex Ratio^{-1}	.029	.007	.000	.007	.039
	(.027)	(.004)	(.007)	(.014)	(.024)
Black×Sex Ratio^{-1}	.005	.015	.045	.098	.002
	(.028)	(.005)	(.011)	(.020)	(.029)
Age	.042	.014	.039	.073	-.066
	(.005)	(.002)	(.003)	(.005)	(.026)
Highest Grade Completed	-.273	-.067	-.207	-.475	.012
	(.059)	(.008)	(.029)	(.052)	(.081)
Both Parents Married	-.466	-.118	-.225	-.389	-.012
	(.087)	(.021)	(.037)	(.110)	(.286)
Year Effects	Yes	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No	Yes
R^2	.05	.068	.103	.117	.03
N	13542	13542	13542	13542	13542

Table 8 Estimated Effect of Sex Ratio on Male's Recent Sex Partners (Corrected SE)

Balanced, All Waves, Highest Black Pop. States					
	POLS	.50	.75	.90	Linear FE
Black	-5.618	-2.414	-7.839	-14.980	(dropped)
	(2.964)	(.586)	(1.395)	(2.514)	
Sex Ratio^{-1}	-.014	.007	-.017	-.034	.025
	(.031)	(.003)	(.009)	(.029)	(.026)
Black×Sex Ratio^{-1}	.063	.025	.085	.166	.039
	(.031)	(.006)	(.015)	(.026)	(.033)
Age	.045	.013	.037	.073	-.055
	(.006)	(.002)	(.002)	(.005)	(.029)
Highest Grade Completed	-.295	-.068	-.203	-.468	-.004
	(.068)	(.010)	(.019)	(.064)	(.100)
Both Parents Married	-.330	-.103	-.182	-.323	-.038
	(.101)	(.020)	(.031)	(.114)	(.392)
Year Effects	Yes	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No	Yes
R^2	.05	.065	.103	.118	.03
N	9956	9956	9956	9956	9956

Table 9 Estimated Effect of Sex Ratio on Male's Recent Sex Partners (Corrected SE)

Balanced, All Waves, Highest Black Pop. States Dropping Sexually Volatile (|DRP|>40)					
	POLS	.50	.75	.90	Linear FE
Black	-4.171	-2.425	-7.260	-12.539	(dropped)
	(1.618)	(.771)	(1.039)	(2.595)	
Sex Ratio^{-1}	-.008	.007	-.012	-.020	.001
	(.012)	(.005)	(.009)	(.014)	(.013)
Black×Sex Ratio^{-1}	.047	.025	.078	.140	.036
	(.017)	(.008)	(.011)	(.026)	(.019)
Age	.035	.012	.036	.072	-.073
	(.003)	(.002)	(.002)	(.004)	(.018)
Highest Grade Completed	-.195	-.064	-.188	-.472	.040
	(.038)	(.013)	(.030)	(.052)	(.067)
Both Parents Married	-.262	-.095	-.182	-.310	.038
	(.063)	(.015)	(.037)	(.108)	(.159)
Year Effects	Yes	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No	Yes
R^2	.10	.074	.124	.149	.06
N	9800	9800	9800	9800	9800

Table 10 Estimated Effect of Sex Ratio on Male's Recent Sex Partners (Corrected SE)

Balanced, 1999-2002 Waves, Highest Black Pop. States Dropping Sexually Volatile (|DRP|>40)					
	POLS	.50	.75	.90	Linear FE
Black	-2.295	.453	.596	-7.479	(dropped)
	(2.416)	(1.340)	(2.566)	(5.595)	
Sex Ratio^{-1}	-.027	.003	-.032	-.026	-.019
	(.020)	(.011)	(.022)	(.034)	(.023)
Black×Sex Ratio^{-1}	.033	.000	.010	.096	.018
	(.025)	(.013)	(.026)	(.055)	(.031)
Age	.026	.015	.036	.050	-.154
	(.004)	(.001)	(.004)	(.010)	(.032)
Highest Grade Completed	-.202	-.039	-.211	-.503	.241
	(.043)	(.023)	(.047)	(.128)	(.095)
Both Parents Married	-.460	-.354	-.498	-.722	-.072
	(.091)	(.041)	(.114)	(.142)	(.213)
Year Effects	Yes	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No	Yes
R^2	.05	.045	.057	.064	.02
N	7741	7741	7741	7741	7741

Table 11 Summary Statistics for Contraception Covariates

Sexually Active, All Waves, Highest Black Pop. States Dropping Sexually Volatile (|DRP|>40)						
	Mean	SD	.10	.25	.50	N
Condom-use Rate	69.303	39.928	0	33.333	100	3549
Age	213.955	28.349	176	194	215	9788
Highest Grade Completed	10.294	2.169	7	9	10	9788
Both Parents Married	.473	.499	0	0	0	9755
Marital Status	.014	.116	0	0	0	9772
4+ partners$_{t-1}$	.226	.418	0	0	1	9788

Table 12 Effect of Race and other Covariates on Contraceptive Behavior (Corrected SE)

Sexually Active, All Waves, Highest Black Pop. States Dropping Sexually Volatile (|DRP|>40)				
	POLS	.10	.25	.50
Black	14.492	5.114	36.937	5.926
	(1.446)	(2.173)	(4.449)	(1.576)
Age	-.490	-.091	-.945	-.292
	(.054)	(.063)	(.096)	(.070)
Highest Grade Completed	2.616	.545	4.713	1.417
	(.563)	(.442)	(1.396)	(.464)
Both Parents Married	4.441	.659	7.862	3.085
	(1.401)	(1.164)	(3.299)	(.838)
Marital Status	-34.614	-1.273	-19.076	-79.995
	(3.595)	(1.650)	(5.868)	(4.475)
4+ Partners$_{t-1}$	.581	1.091	3.053	-.627
	(1.820)	(1.556)	(3.156)	(1.077)
Year Effects	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No
R^2	.12	.007	.135	.074
N	3524	3524	3524	3524

Table 13 Effect of Sex Ratio on Contraceptive Behavior (Corrected SE)

Sexually Active, All Waves, Highest Black Pop. States Dropping Sexually Volatile (|DRP|>40)					
	POLS	.10	.25	.50	Linear FE
Black	-3.591	175.740	104.851	-91.424	.000
	(34.555)	(55.629)	(73.131)	(46.720)	(.000)
Sex Ratio^{-1}	.058	-.000	.541	-.427	.753
	(.346)	(.230)	(.503)	(.515)	(.581)
Black×Sex Ratio^{-1}	.159	-1.499	-.689	.968	-.328
	(.354)	(.479)	(.709)	(.496)	(.590)
Age	-.513	.000	-.923	-.340	.921
	(.059)	(.040)	(.140)	(.075)	(.553)
Highest Grade Completed	2.628	.000	4.571	1.215	-.050
	(.564)	(.239)	(1.065)	(.631)	(1.213)
Both Parents Married	4.431	.000	7.182	2.376	-.486
	(1.401)	(.684)	(3.307)	(.975)	(3.368)
Marital Status	-34.272	-2.000	-19.575	-76.939	-17.681
	(3.620)	(2.651)	(3.620)	(2.753)	(6.382)
4+ Partners$_{t-1}$	.525	-.000	2.248	-1.053	.506
	(1.825)	(.654)	(4.810)	(.864)	(1.950)
Year Effects	Yes	Yes	Yes	Yes	Yes
State Effects	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effects	No	No	No	No	Yes
R^2	.12	.014	.136	.078	.09
N	3524	3524	3524	3524	3524