Increases in Sex with Same-Sex Partners and Bisexual Identity Across Cohorts of Women (but Not Men)

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Abstract

We use data from the 2002-2013 National Surveys of Family Growth to examine change across U.S. cohorts born between 1966 and 1995 in whether individuals have had sex with same-sex partners only, or with both men and women, and in their identity as bisexual or gay. Adjusted for age, race/ethnicity, immigrant status, and mother’s education, we find increases across cohorts in the proportion of women who report a bisexual identity, who report ever having had sex with both sexes, or who report having had sex with women only. By contrast, we find no cohort trend for men; roughly 5% of men in every cohort have ever had sex with a man, and the proportion claiming a gay or bisexual attraction also changed little. We offer an explanation of the gender difference in trends that highlights how the gender system intersects with heterosexism such that the gender revolution and the rising acceptance of gay rights increased the acceptability of having same-sex partners for women more than men.

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Bias against those who are (or appear to be) lesbian, gay, or bisexual is still common in the U.S. (O’Brien 2001; Pascoe 2007; Herek 2008; Mishel 2016; Tilcsik 2011). Nonetheless, change is apparent: a gay rights movement is now decades old (Clendinen and Nagourney 2001), public opinion has moved steadily in a more tolerant direction since the 1990s (Lewis 2015; Ford and England 2016), and there are visible “gayborhoods” in many cities (Ghaziani 2014). In the summer of 2015 a landmark Supreme Court decision said that same-sex marriage must be legal throughout the land (Obergefell v. Hodges 2015).

Despite all this change, we lack systematic analysis of whether there has been change across birth cohorts in the proportion of individuals having same-sex partners or identifying as bisexual, gay, or lesbian. To answer this question, we take a cohort perspective, and use data on adults 18-45 years of age from the National Surveys of Family Growth (NSFG) collected between 2002 and 2013. Separately for men and women, we examine change across cohorts born between 1966 and 1995, using models that adjust for respondents’ age. We do not adjust for survey year, so our measures of net cohort change may encompass cohort and period effects. We also examine whether trends have been significantly different for whites than for blacks, U.S.-born Hispanics,
and Hispanic immigrants. We also ask whether trends differ by socioeconomic background, measured with mother’s education. To foreshadow, we find substantial evidence of increases in women’s sexual behavior with same-sex partners, and in bisexual identity, but little change for men. Trends differ little by race, ethnicity, immigration status, or socioeconomic background. We conclude by offering a speculative explanation of why change occurred for women, but not men.

**PAST RESEARCH ON CHANGE IN SEX WITH SAME-SEX PARTNERS**

**Change Across Periods.** Anderson and Stall (2002) found an increase from 1-2% to 3-4% between 1988 to 2000 in the percent of men who had had sex with a man in the last year, but no significant increase in the proportion having *ever* had sex with a man. They used data from the General Social Survey (hereafter GSS), and did not examine trends for women. Turner et al. (2005) used the same 1988-2002 GSS data and found substantial increases in same-sex sexual behavior for women in the 1990s in whether women had sex with a same-sex partner in the last year, the last 5 years, or ever. They found increases for men in same-sex activity in the last year, but they were much smaller than the increases for women.

Butler (2005) also used the 1988-2002 GSS data, augmented by the 1992 National Health and Social Life Survey, to examine period change. She found increases in the percent of men and women who had sex with a same-sex partner in the previous year across the period. Using a linear functional form for period, the rate of change was significantly larger for women than men.

A more recent analysis by Copen et al. (2016), using a subset of the NSFG data we use, found an increase between the 2006-10 and the 20011-13 NSFG surveys in the
proportion of women who reported same-sex contact, and in the proportion of both men and women who claimed a bisexual identity.

Thus, taken as a whole, these studies, all of which focused on period trends centered on the 1990s, suggest increases for both sexes in sex with same-sex partners, with larger changes for women. While our focus is on the U.S., we note that Johnson et al. (2001:1839, cited in Turner et al. 2005:459) found increases for British men and women in same-sex contact during the preceding five years during the 1990s.

Change Across Cohorts. Butler (2005), whose main analysis focused on period change, also provided descriptive information on change in having had a same-sex partner since age 18 across cohorts born between 1929 and 1982. She found significant increases for women, but not for men. These analyses, however, had no control for age, or other compositional covariates, and thus could be a function of period change or compositional demographic change. Turner et al. (2005), mentioned above for their period analysis, also examined the proportions of men and women reporting any same-sex partner since they were 18, and how it varied by cohort. For women, they found increases from 1.6% for the cohort born before 1920 to 6.9% for those born between 1970 and 1984, the most recent cohort they examined. They found no cohort change for men. This analysis was from a model with no covariates, even for age; by contrast, their period models contained covariates. Thus, we have no multivariate cohort-focused analysis of change in sexual orientation or sex with same-sex partners.

Gender Differences in Changes. While Turner et al (2005), Butler (2005), Copen et al. (2016), and Johnson et al. (2001, using British data) all found increases for both men and women in sex with same-sex partners, Turner et al. (2005) and Butler
THEORIZING WHY CHANGES IN SEX WITH SAME-SEX PARTNERS DIFFERS BY GENDER

Nature, Nurture, and Change. Explanations of sexual behaviour or orientation often involve questions about the role of biology, social forces, and their interaction. The hypothesis we will put forward below about the change we find in women’s, but not men’s, sexual orientation and behavior is social. It is consistent with a completely socially constructionist notion of what explains sexual orientation and behavior. It is also equally consistent with a view that sees strong genetic effects on these phenomena, provided that biology is not seen to explain all the variation, leaving nothing to be explained socially. (For evidence regarding genetic effects on sexual orientation, see Bailey and Pillard 1991; Bailey et al. 1993; Bailey et al. 2000. For a critical review Bearman and Brückner 2002.) Our conceptual work is not directed at the basic question of why people have sex with men, women, or both, but why there has been change across cohorts in these sexual behaviours and associated identities, and why the change might have differed for men and women.

Increased Cultural Acceptance of Sex with Same-Sex Partners. The simplest, most obvious, and gender-neutral hypothesis about why we would see increased same-sex sexuality is that social norms and institutional rules became more accepting. Butler (2005) suggests this. She shows evidence of steady increases since the 1970s, escalating in the 1990s, of states and localities repealing sodomy laws and passing laws against discrimination based on sexual orientation in the workplace. She also notes the rise in employers offering benefits to same-sex domestic partners. Such change undoubtedly
reflected as well as reinforced change in social norms. To examine change in norms, she
uses GSS attitude data on views of whether same-sex sexual relations are wrong
(measured on a 4-point scale ranging from “always wrong” to “not wrong at all”). As far
as we know, no one has examined cohort change in these views, but Butler finds
substantial liberalization of attitudes across the period of the 1990s (Butler 2005:426).
Others have found this as well (Anderson and Fetner 2008; Lewis 2015; Ford and
England 2016).1

In an attempt to test whether normative change influenced the behavioral change,
Butler (2005) constructed year- and region-specific averages on the attitudinal measure
discussed above, and entered the variable in her models predicting whether the
respondent had engaged in sex with a same-sex partner in the last year. Using separate
regressions for men and women, she finds that the coefficient for period (specified in
linear fashion) was no longer significant in the model for men after entering the
attitudinal control, but entering the control caused only a small (16%) reduction in the
period coefficient for women, which remained significant (Butler 2015:441). This
suggests that changing attitudes led to the increase in same-sex sex for men, but didn’t
explain the larger increase for women. However, as Butler points out, we cannot be sure
that the significant coefficient on the attitudinal measure is indicative of a causal effect of
the norms on the behavior. It could be that an increase in sex with same-sex partners is
what changed the norms, rather than the reverse, or that both were caused by a third
factor.

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1 Whether change in liberalization of attitudes started earlier is unclear. Lewis (2015) shows some
liberalization in the 1970s, reversing in the 1980s. Thereafter, he shows liberalizing attitudes, as do other
studies.
Theorizing Gender-specific Change. The studies we have reviewed suggest significant recent change among women in same-sex sexuality. However, depending on the study or whether cohort or period change was examined, little or no change was found among men. Why might change occur among women, but much less, if at all, among men? An answer is revealed if we focus on how the gender system intersects with heteronormativity or heterosexism. In particular, what we need to understand is how norms about sexuality with same-sex partners are linked to norms about gender.

We start by pointing out two distinct aspects of the gender system. The first is that people face social pressure to conform to what is expected of them as men or women. Gender conformity entails many things, such as that men should be strong and women nice. The gender system is relevant to our inquiry about sexuality because conforming to gender norms also requires at least appearing to be heterosexual. That is, just as women violate gender norms by being executives or carpenters and men violate gender norms by being stay-at-home dads or nurses, both men and women violate gender norms by not appearing to be straight (Watts 2015; Pascoe 2007). Violating gender norms is seen as negative in many quarters and thus often has social costs.

A second aspect of the gender system is that activities or attributes associated with women tend to be valued less than those that are associated with men. As one example of this, jobs filled largely by women pay less than what you could expect them to pay based on their demands (England 1992; Kilbourne et al 1994; Levanon and England 2009; for debate see Tam 1997, 2000; England et al. 2000). As another example, being a soldier is associated with men and parenting with women. The state provides a
special health care plan for veterans of military service, but there is no analogue for mothers.

When we put these two aspects of the gender system together—that people face disapproval for violating gender norms, and that masculinity is more highly valued than femininity—it implies that men will lose more status by violating gender expectations than women will. When men violate gender norms, they are seen as more feminine; when women violate gender norms, they are seen as more masculine. Either may be met with social disapproval, but we believe that the stigma is greater for men’s gender nonconformity precisely because it is seen as a movement toward a devalued status—femininity. It is consistent with our argument that boys seen as effeminate are stigmatized much more than girls seen as boyish (Fine 1987; Pascoe 2007). When we add to this that being straight is seen as nonconformity to gender norms, it implies that departures from exclusive heterosexuality are more stigmatizing for men than women. Experimental work by Watts (2015) has shown exactly this. Other research supports the conclusion as well; ridiculing someone by calling him a “fag” is common among male youth, but it is much less common to ridicule girls by calling them “dykes” (Pascoe 2007). The greatest stigma seems reserved for men who appear effeminate and identify as gay or engage in sex with same-sex partners, especially those being the receptive partner in anal sex.²

Given this gender difference in the stigma associated with violations of gender norms, we would expect women to have always engaged in gender deviations more than men, because penalties for men are so much greater. That doesn’t necessarily suggest a

² A clear exception to this generalization is the set of ritualistic sexual practices between men who identify as straight discussed by Ward (2015). Yet, the fact that the men Ward studied shun a gay or bisexual identity strongly suggests that men are stigmatized more than women for sex with same-sex partners or non-heterosexual identities, as experimental work by Watts (2015) has shown.
stronger upward trend for women in sexual or other behavior that challenges gender norms, but suggests a higher base level all along. Yet, as an empirical matter, what we call “the gender revolution” clearly has entailed more change in women’s activities than men’s (England 2010). Many women enter male professions; fewer men enter female jobs or become full-time homemakers. Girls now play sports, while fewer boys play with dolls. Women wear pants, but men still eschew wearing dresses. Moreover, we get used to seeing women do formerly transgressive things to the point that they are normalized, hardly registering them as gender violations.

Why has change in gendered activities of all sorts been greater for women than men? Women had economic incentives to seek employment and enter better paying, male-dominated jobs, while men had no economic incentive to become stay-at-home dads or devote more time to child rearing. Once discriminatory barriers were reduced, these greater incentives were undoubtedly important in generating change. But economic incentives would hardly seem to apply to women’s changing sexual behavior; there is no reason to think women have an economic incentive to seek same-sex sexual partnerships.

We suggest the following explanation of why there has been more increase in women’s than men’s departure from exclusive heterosexuality. Changes associated with the gender revolution sent the message that gender nonconformity was more acceptable than before, and, given that deviations from exclusive heterosexuality are seen as gender nonconforming, part of the implicit message was permission to have same-sex sexual relationships. The rise in norms accepting of gay rights had the same effect. In short, our argument, albeit speculative, is that both change in gender norms and change in attitudes about gay rights had the potential to reduce the social costs of deviating from exclusive
heterosexuality and encourage same-sex sexual relationships for women. But the way that cultural conceptions and norms about gender and sexuality are intertwined, combined with the negative view of anything coded as feminine, conspired to encourage less increase in same-sex sexual relationships for men.

LIMITATIONS OF PAST RESEARCH AND OUR CONTRIBUTION

Past research suggests an increase in sex with same-sex partners, but none of these studies distinguished between those who have sex with both men and women and those who have sex only with same-sex partners (either ever, or in a given time period). Our analysis will make this distinction, and we will show that it is important to understanding the trends among women. Past research is also limited in that we found no studies other than Copen et al. (2016) that examined period or cohort change in reported sexual orientation, and they focus on change across a very short time period (2006-10 to 2011-13). Finally, no one has performed a systematic multivariate investigation of cohort change; each of the only two analyses of cohort change, by Butler (2005) and Turner et al. (2005), were bivariate comparisons without covariates. Additionally, our analysis has data on more recent cohorts, including those born in 1985-95. In short, our empirical contribution is to provide the first multivariate analysis of cohort change in having sex with same-sex partners only, having sex with same- and other-sex partners, and identification as bisexual or gay. Our conceptual contribution is to discuss how gender and heterosexism are intertwined, and, from this, provide an explanation of the increase in women’s - but not men’s - departure from exclusive heterosexuality.

DATA AND METHODS

Data
We pooled data from the 2002, 2006-2010, and 2011-2013 waves of the National Survey of Family Growth (NSFG). In all of these waves, there was an Audio Computer-Assisted Self-Interview (ACASI) section that included questions on whether one had same-sex sexual partners ever, and sexual orientation. Since this portion of the questionnaire was not face-to-face (the interviewer let the respondent answer on the computer), answers about stigmatized behaviors and identities are expected to have less reporting bias. (For evidence that bias is reduced with ACASI, see Villarroel et al. 2006). Of course, it is still entirely possible that nonheterosexual behaviors and identities are underreported.

Variables

**Dependent Variables: Sexual Behavior.** To construct our behavioral dependent variables regarding sexual behavior, we used the questions on same-sex sexual partners in the ACASI section, and the questions on sexual intercourse with other-sex partners in the main questionnaires to identify whether each respondent had: 1) had sex with other- and same-sex partners, 2) had sex only with same-sex partners. (Those not in the preceding categories of interest had had sex only with other-sex partners, or with no one.) Respondents were asked questions that put them in these categories with respect to their sexual behavior ever, and their behavior in the last 12 months.

The specific questions we used from the ACASI section are: “Thinking about your entire life, how many [same-] sex partners have you had?” and “Thinking about the last 12 months, how many [same-] sex partners have you had?” Women were asked these questions if they stated yes to a prior question about whether they had ever had “sexual experience of any kind with another woman” or yes to one of two prior questions about
whether they had ever given or received oral sex to/from a woman. (In 2002 only, the oral sex questions were not asked, so women were asked how many same-sex partners they have had, last year or ever, only if they answered yes to ever having had sexual experience with a woman.) Men were asked their number of male sexual partners, ever or last year, only if they answered yes to either of two previous questions regarding whether they had ever had oral or anal sex with a man. Respondents who did not get asked the questions about number of same-sex partners because of their negative responses to the prior questions were assumed to have had no same-sex partners. To measure other-sex sexual activity, we relied on the NSFG official recodes for number of other-sex sexual partners with whom respondents had intercourse ever, and in the last 12 months, based on information provided in face-to-face interview using the survey’s main questionnaires. Using the combination of these variables allowed us to ascertain whether each respondent had ever had sex with at least one man and woman, or only with one or more women, or one or more men. This was our main behavioral dependent variable. In a supplementary analysis, we also used these measures of behavior for the last year.

One potential problem in comparing women’s and men’s levels of sex with same-sex partners is that the different screening questions used to path men and women into questions about same-sex partnerships appear at first glance to have created a higher bar for men than for women to be seen as having had a same-sex partner. That is, men weren’t asked how many male sexual partners they had had (and thus were assumed to have had none) unless they said they had oral or anal sex with a man, whereas women could be classified as having had sex with a woman if they said they had had any sexual experience with a woman, even if they did not report having had oral sex with a woman.
Recent attention to the prevalence of women kissing women on dance floors and at parties (Rupp et al. 2014; Hamilton 2007) raises the question of whether women reporting sexual experience with women are referring to experiences such as these or to more private sexual contact involving genitals. In analyses not shown we ascertained that 91% of women of age 18-45 who said they had sex with a woman last year (regardless of whether they also said they had sex with a man) also reported that they had ever had oral sex with a woman, as did 88% of women who reported having sex with both men and women last year.3 This suggests that the vast majority of those who say they have had a female sexual partner have had private sexual experiences with women beyond kissing. This makes us relatively unconcerned that the measures artifactually create a higher bar for men than women reporting on sexual behavior.

**Dependent Variables: Identity and Attraction.** To assess identity, we used a question that asked respondents whether they see themselves as “heterosexual or straight,” “homosexual, gay, or lesbian,” or “bisexual.”4 Whether we are referring to men or women, we will use the term “gay” (as opposed to “lesbian”) for brevity and to facilitate comparing analyses across genders. One complication is that in the 2002 wave and part of the 2006-2010 wave (through June 2008), respondents were given the option of choosing “something else” to describe their sexual orientation; we treated respondents who gave this response as in the reference category (with heterosexuals).

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3 By contrast, only 71% of women 18-45 years of age who called themselves bisexual report ever having had oral sex with a woman. Consistent with this, using these same data, Brown and England (2016) found that 23% of bisexual women have never had sex with a woman (16% have had only male partners, and 7% no partners of either sex).

4 The exact words used for the “gay” option varied by sex and year. In 2002, the option was “homosexual” for both men and women. In 2006 and later, the “gay” option for women was “homosexual, gay, or lesbian” while for men it was “homosexual or gay.”
In supplemental analyses we also use attraction, a 5-point scale with categories for only attracted to females, mostly attracted to females, equally attracted to males and females, mostly attracted to males, and only attracted to males (with gender order reversed for female respondents). We treated respondents who chose “not sure” as having a missing value, so they were eliminated from the analysis. We recoded the categories into heterosexual attraction (only attracted to the other sex), bisexual attraction (attracted mostly to either sex, or equally to both), and gay attraction.

**Independent Variables.** Our main independent variable of interest is respondent’s birth cohort, represented by indicator variables. The categories containing enough respondents for analysis are 1966-74, 1975-79, 1980-84, and 1985-95; we eliminated respondents born between 1957 and 1965. We do not enter period into our main models, but we do employ detailed controls for age—indicator variables for each single year of age at survey.

Control variables include indicator variables for the following race/ethnic groups: Non-Hispanic White (hereafter white), Non-Hispanic Black (hereafter black), Hispanic, and Other. We also control for immigration status (=1 if not born in the US), and an interaction for Hispanic X Immigrant, that preliminary analysis showed was needed. The other control we used was respondent’s mother’s education, represented by indicator variables for completing Less than High School, High School (including those with some college but not a Bachelor’s degree), or a Bachelor’s degree or higher.

Means on all variables, separately for men and women, are in Table 1. All of our analyses incorporate survey weights corresponding to the relevant data collection years.

**Models**
**Main Models.** We estimated a series of logistic regressions. Each regression is comprised of the cohort indicators as the primary independent variables, as well as all the control variables discussed above. Our goal is to get point estimates of the amount of change in each dependent variable between cohorts. The controls for age are intended to help us distinguish between cohort and life cycle differences, given that the cohorts differ in the age distribution of respondents in the survey. Given the familiar age-period-cohort identification problem, we do not put the survey year (period) in our main models, and thus the cohort *trends* that we estimate may contain a combination of cohort and period effects. We control for race, ethnicity, immigration status, and mother’s education because these variables may affect sexual behavior or identity, and, thus, compositional change across birth cohorts might create cohort effects. For example, we will find a much lower rate of same-sex sexual behavior by Hispanic immigrants than other groups; given this, if the proportion of successive cohorts made up of Hispanic immigrants changed, failure to control for this compositional factor would, all else equal, lead cohorts to differ in sexual behaviors. We seek to identify change across cohorts that is not driven merely by compositional changes but by changing behavior within groups, and the controls described above accomplish this.

All models are estimated separately for men and women, allowing us to examine whether trends differ by gender. For each gender, our main models entail estimating logistic regression models predicting the following dependent variables:

- **Model 1:** Has had male and female sexual partners ever
- **Model 2:** Has had only same-sex sexual partners ever
Each of these has a reference category that contains all other categories. Thus, the reference categories for Models 1 and 2 are overlapping such that those who have had no sexual partners or only those of the other sex are in the reference category for both models, while those with both sexes as partners are in the reference category only for Model 2 and those with only same-sex partners are in the reference category only for Model 1.

To examine cohort change in sexual orientation (identity), we estimate two logistic regression models that predict dependent variables measuring the sexual orientation the respondent identifies with:

Model 3: Bisexual

Model 4: Gay

Each of these has a reference category that includes the other, and also includes those who answered heterosexual (and, in the years when it was an option, something else).5

Supplementary Analyses. In a supplementary analysis discussed but not displayed in our regression tables, we estimated models parallel to 1 and 2, regarding sex with same- and other sex partners, or same-sex partners only, but referring to partners in the last year (rather than ever). In another supplementary analysis discussed but displayed in our regression tables, we ran models parallel to 3) and 4) but taking bisexual and gay attraction (as specified above) as the dependent variables, rather than identifying as bisexual or gay.

5 An alternative would be to estimate one multinomial logistic regression instead of Models 1 and 2, where the two competing choices are sex with same- and other-sex partners and sex with only same-sex partners, each relative to a reference of those who had no or only other-sex partners. Analogously, instead of Models 3 and 4, we could have estimated one multinomial logistic regression. When we do this for the whole sample, this gives similar results to those we report. However, because, for supplementary analyses, we estimate separate models for whites, blacks, U.S.-born Hispanics, and Hispanic immigrants, the MNL approach leads to coefficients that cannot be estimated because of empty cells.
We also undertake supplementary analyses to examine whether cohort trends we identify in any of our main models discussed above differ by race, ethnicity, or nativity. To do this, we estimate separate models for whites, blacks, U.S.-born Hispanics, Hispanic immigrants, and others. We then examine the significance of differences between cohort coefficients in the models for whites and each other group. We discuss these findings, but they are not shown in our regression tables. Analogously, to see if those from different socioeconomic backgrounds experienced distinct trends, we estimate separate models for each group by mother’s education, and examine the significance of differences between cohort coefficients for those whose mothers were at least college graduates and each other group.

**Predicted Probabilities.** From all of our models, we calculated predicted probabilities of observing the sexual behaviors, identities, or attractions of interest across cohorts.

**Sensitivity Tests.** We also undertook sensitivity tests to examine the robustness of our conclusions about cohort differences. First, we estimated the hierarchical age-period-cohort model proposed by Yang and colleagues (Frenk, Yang and Land 2013; Yang and Land 2006; Yang and Land 2008). In the original two-level model, age is included as a covariate in the first-level equation, which explains variation at the individual level, and cohort and period are regarded as contextual effects, and are assumed to be contributions to a random intercept in the second-level equation. This ensures that there is no perfect dependence between these variables, because the parameters for cohort and period are not directly estimated (Yang and Land 2006). Bell and Jones (2014) examined the performance of the Yang hierarchical model using
simulated data and concluded that this model was unable to tell apart period from cohort effects in the presence of cohort effects. When cohort trends are suspected, they recommend adding cohort as a fixed effect in the first-level equation to absorb the systematic variation in cohort effects, while still specifying period variation as a random intercept. This hierarchical random intercepts model decomposes the residual variance, which remains after controlling for age and cohort, into period-related variance, and individual, idiosyncratic variance. This way of estimating the size of cohort effects entails making the assumption that, after controlling for age and cohort, period contributions represent a random disturbance uncorrelated with the idiosyncratic error and uncorrelated with the covariates in the first-level equation (Rabe-Hesketh and Skrondal 2012), clearly a strong assumption. It gains some plausibility from the fact that the 12-year period range that we observe (2002-2013) is much shorter than the 30-year birth cohort range and the 45-year age window, along which it is more feasible to find systematic effects. Our hierarchical model accounts for the clustering in respondents across periods, and adjusts standard errors accordingly, in this way making the model a more conservative analysis than our main regressions, which did not account for the period clustering of the data. These regressions included the same controls as the models in our main analysis.6

Our more conservative sensitivity test was to add indicator variables for periods to a variant of our original models that control for age and cohort. In order to preserve statistical power and avoid small-cell problems, instead of the detailed age indicator variables, we substitute three terms, for age, its square, and its cube to capture age effects more parsimoniously, but still allow them to be nonlinear. To account for period, we

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6 Our hierarchical models are estimated using GLLAMM in Stata 13, and follow the recommendations given by Rabe-Hesketh and Skrondal (2006) to rescale survey weights.
include three dummy variables, identifying interview years 2002, 2006-2009, and 2010-2013. Other than age, and the addition of period, variables are as in our main regressions. These variable definitions avoid the perfect linear combination problem that would arise if cohort, age, and period were introduced as linear terms. Because age is highly correlated with cohort, and cohort is correlated with period, including controls for all three of them is a very conservative strategy. However, these and the hierarchical models still rely heavily on the assumption that the interval width of cohort and period variables correctly captures cohort and period effects, and that age is adequately controlled for by high-order polynomials (Luo and Hodges 2015).

RESULTS: CHANGES IN WOMEN’S BEHAVIOR AND IDENTITY

Overall Trends

Table 2 presents our main regression analyses for women, where cohort change is estimated under controls as detailed above. Model 1 reveals the cohort change in whether women have had sexual partners of both sexes (ever), while Model 2 does the same for having had only women as partners. Each successive cohort shows a significant elevation from the reference birth cohort (1966-74), and increases are monotonic. Model 3 also shows significant monotonic upward trends in identifying as bisexual. Model 4 shows some increase in identifying as gay in the last two cohorts, but not enough to be statistically significant.

We can see the levels of these behaviors and identities and the magnitude of change more clearly in Table 3, containing predicted probabilities for each cohort, calculated from the models in Table 2 (and some supplementary models on related outcomes). Table 3 shows that the predicted percent of women who had ever had sex
with both men and women increased from 10.3% in the 1966-74 birth cohort, to 19.7% in the 1985-95 cohort. The percent of women who had ever had sex with only female partners, always very small, nonetheless rose significantly from 0.2% to 1.5% from the first to last birth cohort. Having had partners of both sexes, as well as having had only female partners in the last year also increased, but the regressions (not shown) don’t show these increases to be significant. Bisexual identity increased significantly from 2.7% to 7.2% across the four cohorts, while bisexual attraction increased significantly from 13.8% to 19.5%. Gay identity, always much smaller than bisexual identity, did not rise significantly, while gay attraction, also much rarer than bisexual attraction, rose at marginal significance.

In sum, across the four cohorts spanning those born from 1966 to 1995, the headline is that having had sex with same sex partners, bisexual identity, and reported attraction to same-sex partners went up significantly among women. The increases for behavior and identity by cohort can be seen clearly in Figures 1 and 2.

**Do Trends Differ by Race, Ethnicity, and Immigrant Status?**

The same general picture of increases across cohorts in sexual experience with same-sex partners applies to all groups of women. To ascertain this, for the four outcomes in Table 2, we estimated separate regressions for the following groups: Non-Hispanic whites, Non-Hispanic blacks, U.S. born Hispanics, and Hispanic immigrants. (Regression results not shown. We do not discuss results for the residual category of other races or non-Hispanic immigrants.) Trends in having ever had sex with women only aren’t different between whites and any other group besides U.S.-born Hispanics, whose rise was significantly faster. Trends in having ever had sex with both sexes are not
significantly different between groups, except that increases for black women are much steeper than those for whites, starting at a (non-significantly) lower level in the first cohort and rising to a much higher level. The other difference is that the upward trend in bisexual identity is not seen for Hispanic immigrants. Predicted values show that Hispanics, especially Hispanic immigrants, have the lowest level of any of these behaviors or identities that indicate lack of exclusive heterosexuality. Overall, the conclusion is that all of these groups have upward trends in sex with same-sex partners that are the same as or steeper than those for non-Hispanic whites. Other than Hispanic immigrants, whose bisexual identity did not rise across cohorts, other groups had a rise in bisexual identity no different than that for white women. Figures 3 and 4 display the cohort trends on same-sex sexual behavior and gay or bisexual identity separately by race.

We also examined whether cohort trends vary by socioeconomic background, as measured by mother’s education, finding almost no differences. Specifically, there were no significant differences between the cohort effects for these groups of women on having had sex with both sexes, having had sex only with women, or gay identity. The upward cohort effects on bisexual identity were larger for the group with the lowest mother’s education, and one of these differences (the elevation between the first and second of the four cohorts) was significant. Overall, the trends occurred among women of all socioeconomic backgrounds.

Our sense is that many people have the perception that the change toward queer behaviors and identities is mainly occurring among socioeconomically privileged white
women, especially those who go to college. Yet, our analysis suggests that these trends largely transcend race and class, and if anything, are steeper in less privileged groups.

**Are Cohort Trend Estimates Robust to Sensitivity Checks?**

As described above, our first sensitivity check is to use a variation of the method proposed by Yang and colleagues (Frenk, Yang and Land 2013). We estimated hierarchical models with a first stage containing cohort as a fixed effect and (linear, squared, and cubic versions of) age, while specifying period variation as random intercepts for individual years. Cohort is coded as in our main models in Table 2. Numerical results are not shown, but we describe them here. Table 2 showed rising trends of having had sex (ever) with both men and women, with odds ratios (ORs) of 1.47, 1.73, and 2.18 for respective cohorts (all significantly different from the reference first cohort). The analogous ORs from this sensitivity test are 1.56, 1.76, and 1.88 (all significant).

Thus, this trend is quite similarly estimated by the two techniques. Sex with same-sex partners only had ORs of 4.01, 7.29, and 8.76 in Table 2 (all significant), and in this sensitivity test, results were 2.06, 3.43, and 4.56 (the last significant); the sensitivity test effects are smaller but still show an upward trend. Similarly, the ORs for successive cohorts predicting bisexual identification, which were 1.67, 1.95, and 2.85 in Table 2 (all significant) also show monotonic increases, with very similar ORs of 1.58, 1.95, and 2.63 (all significant). Gay identity did not increase at conventional significance levels in either analysis. Thus, this alternative and more conservative method gives qualitatively similar results. This increases our confidence in our conclusion of cohort change, although it does not do much to increase our confidence that the cohort trends we

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7 This conclusion is bolstered by the finding that the intra-class correlation coefficients, which represent the proportion of the residual variance that is attributable to period differences is very low (less than 1%).
identified are indeed cohort, rather than period, effects, since the assumption of the model is that period effects are random and uncorrelated with cohort effects.

Our second sensitivity test is aimed at discerning whether the coefficients in our main models represent cohort rather than period effects. As described above, we do this by entering age, period, and cohort in the model in such a way that the model is still identified. Cohort enters as in our main regressions in Table 2, age enters here as linear, squared, and cubic terms, and period enters as two dummy variables for the three categories, 2002 (the reference), 2006-08, and 2010-13. In these results not shown, we find increases across cohorts in having had sex with both sexes that are positive and (for the first two of the three cohort dummies) significant, although they are smaller in magnitude than the significant increases in our original analysis in Table 2. Sex with only women trended significantly upward in Table 2, but was non-significant in this more conservative analysis. This analysis also shows increases in bisexual identification, but they are not significant, whereas they were significant in the prior analysis in Table 2. Gay identity showed no significant trend in either analysis. Our conclusion is that findings in this more conservative analysis are also suggestive of upward cohort effects, most clearly for having had sex with both women and men. But the evidence for increases in sex with women only, and in bisexual identity found in our original analysis does not show up as significant in this more conservative approach. In addition, both of our sensitivity test strategies rely heavily on the assumption that the coding of cohort, age, and period correctly captures their effects. Thus, while this gives us some confidence that there are some true cohort effects in having had sex with both sexes, overall, we
think it safer to conclude that we have found cohort trends, and not make strong
statements about cohort effects for the results writ large.

RESULTS FOR MEN

We find no upward trend across cohorts in whether men have had sex with only
men, or with both men and women, either ever or last year, and no upward trend in either
gay or bisexual identity or attraction. The last cohort is never significantly above the first
in any of the outcomes (Tables 4 and 5). On two outcomes (partners of both sexes ever
and gay identity) the likelihood went down between the first and second cohort. This lack
of upward trend is not significantly different for groups of men from varying class
backgrounds, defined by their mother’s education (results not shown). The lack of trend
is true within race/ethnic and nativity groups as well for most outcomes (regression
results not shown, but see Figures 4 and 5 for cohort trends on same-sex sexual behavior
and gay or bisexual identity, separately by race).8

Our findings differ by gender, not only in the lack of an upward trend for men, but
also in the lower baseline levels of departures from exclusive heterosexuality for men, as
can be seen in Figures 1 and 2. As Table 5, giving predicted probabilities, shows, in every
cohort, between 3.5% and 4.8% of men have ever had both sexes as partners, and
between 0.8% and 1.5% had had only same-sex partners, with no significant upward
trend in either. Adding the two groups together, between 4% and 6% had ever had a
same-sex partner in each cohort; by contrast, for women this was 10.5% in the first

8 There are no significant differences in cohort coefficients between whites and blacks, U.S.-born
Hispanics, or Hispanic immigrants on sex with other and same-sex partners ever, sex with only same-sex
partners ever, or bisexual identity. The only exception for the outcomes in Table 4 is that U.S.-born
Hispanics had a significant increase in identifying as gay the last compared to first cohort; no such increase
was present for whites, and the differences between the two group’s coefficients for the last cohort are
significantly different.
cohort and 21.2% in the last (Table 3). The percent of men identifying as gay was 2.8%, 1.4%, 1.6%, and 1.7% respectively in the four cohorts, with the percent of those claiming to be bisexual rising from 1.3% to 2.4%. None of these indicators shows significantly higher proportions in the last than first cohort for men. (Nor were trends in other indicators in Table 5 significant for men.) When we estimate the same alternative models for men that we presented for women as sensitivity tests, they also suggest a lack of cohort change. Thus, the clear conclusion is that things changed for women, but not for men, and that by the most recent cohort a much higher proportion of women than men had had sex with both sexes, and claimed a bisexual identity.\(^9\)

**DISCUSSION AND CONCLUSION**

We have shown a steady increase across cohorts in women’s experience with same-sex partners. The upward trend is present for women having had sex only with women, as well as having sex with both sexes. The proportion of women identifying as bisexual has also grown. The upward trend in experience with both sexes is no less steep among blacks, U.S.-born Hispanics, and Hispanic immigrants than it is among whites; for a few outcomes increases are steeper for minority groups. Hispanic immigrants differ from other groups of women in being less likely to identify as something other than heterosexual or to have had sex with a same-sex partner.

When we take the conservative approach of entering period in the model along with age and cohort as a sensitivity test, it appears that increases in women having sex

---

\(^9\) This conclusion could be challenged under some assumptions about under-reporting. We might think that due to the greater stigma for men than women of gay sex, that men would under-report more. If the gender gap in under-reporting remained constant, but men always under-reported more, then we might be overestimating the gender gap in behavior and identities, but our estimates of trends in behavior, although not levels, should be accurate. Another possibility is that women’s under-reporting decreased (for some of the reasons of changing permissiveness we discuss) but their behavior did not change. However, we doubt that the more than doubling of having had a same-sex partner and of bisexual identity was all just a change in reporting.
with both men and women are mostly cohort “effects,” but this conservative approach
does not show significant cohort effects on women’s bisexual identity. Thus, we think it
safer to conclude that the cohort change we have identified among women reflects a mix
of cohort and period effects.

By contrast, there is no apparent trend in having only same-sex sexual partners, or
both male and female partners, for men. Nor is there an increase in the proportion who
identify as gay or bisexual. For the most part, these “non-trends” are not significantly
different for whites than for blacks, U.S.-born Hispanics, or Hispanic immigrants, nor do
they differ across groups of men defined by their mother’s education.

Our data are not rich enough to allow us to test various possible explanations of
the cohort trends that we find. What we can offer is a speculative explanation that
comports with the central fact shown by our analysis—that movement away from
exclusively heterosexual behavior and identity has been greater for women than men. We
theorized that two factors each facilitated such change across cohorts of women: the
gender revolution and increased acceptances of gay rights. While the cultural devaluation
of activities traditionally associated with women has changed little, a message of the
gender revolution that did take hold was that it is all right for women to do things only
men could previously do. This increased the sense that it was permissible for women to
engage in sex with women, even though doing so is still a violation of traditional gender
conformity. Increased tolerance for gay rights furthered the sense of permission.

One might have thought that the same two factors, greater acceptance of gay
rights and acceptance of more flexible gender expressions, would have furthered
acceptance of men’s same-sex relationships in a parallel fashion. But we argued that the
way in which the gender system intersects with heterosexism deters acceptance of men’s
defying gender expectations to have sex with men. While some feminists urged a
revalorization of traditionally feminine activities as well as an acceptance of men doing
things seen as feminine, this was certainly not the message the average person received
from the gender revolution, which mostly involved changes in women’s roles. Because
same-sex relationships continued to be seen as gender-bending, and the continued
devaluation of the feminine meant that same-sex relationships entailed losing status for
men more than it did for women, there was less change in the acceptance of men’s
departures from exclusive heterosexuality.

Of course, some men and women have always expressed their same-sex desire,
despite stigma, and we do not want to understate the social costs women still experience
for departures from exclusive heterosexuality. Nonetheless, our conjecture is that changes
in norms about gender and heterosexism intersected in a way that provided more of an
opening for women than men to have same-sex relationships. While our explanation is
speculative, what we have shown clearly is that there has been more increase across
cohorts in reported sex with same-sex partners, and in bisexual identity, for women than
men.
REFERENCES


Figure 1: Women's and men's predicted probability of having had sex with same- and other-sex partners, by cohort
Figure 2: Women's and men's predicted probability of identifying as bisexual or gay, by cohort
Figure 3: Women's predicted probability of having had sex with same- and other-sex partners, by cohort and race
Figure 4: Women's predicted probability of identifying as bisexual or gay, by cohort and race
Figure 5: Men's predicted probability of having had sex with same- and other-sex partners, by cohort and race
Figure 6: Men's predicted probability of identifying as bisexual or gay, by cohort and race
<table>
<thead>
<tr>
<th></th>
<th>Women</th>
<th>Men</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Sexual Behavior</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Same- and other-sex partners</td>
<td>0.136</td>
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</tr>
<tr>
<td>ever</td>
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<tr>
<td>Only same-sex partners ever</td>
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<td>0.011</td>
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<tr>
<td>ever</td>
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<td></td>
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<td>Only other-sex partners ever</td>
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</tr>
<tr>
<td>Born 66-74</td>
<td>0.357</td>
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</tr>
<tr>
<td>Born 75-79</td>
<td>0.201</td>
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</tr>
<tr>
<td>Born 80-84</td>
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</tr>
<tr>
<td>Born 85-95</td>
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<tr>
<td><strong>Race, Ethnicity, and Nativity</strong></td>
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</tr>
<tr>
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<td>0.625</td>
</tr>
<tr>
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<td>Non-Hispanic Immigrant</td>
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<td><strong>Mother's Education</strong></td>
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<td>Less Than High School</td>
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</tr>
<tr>
<td>HS/Some College</td>
<td>0.570</td>
<td>0.578</td>
</tr>
<tr>
<td>BA or more</td>
<td>0.195</td>
<td>0.210</td>
</tr>
</tbody>
</table>

Note: N of entire sample = 22,954 women and 17,452 men; Ns for individual means vary according to missing values for the variable.
Table 2: Odds Ratios from Four Logistic Regressions Predicting Whether Women:
1) Have Had Male and Female Sexual Partners (Ever),
2) Have Had Only Female Sexual Partners (Ever),
3) Identify as Bisexual, and 4) Identify as Gay

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
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<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
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<tr>
<td><strong>Cohort (ref: 66-74)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Born 75-79</td>
<td>1.467**</td>
<td>4.012**</td>
<td>1.673**</td>
<td>0.944</td>
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<tr>
<td></td>
<td>[0.153]</td>
<td>[1.860]</td>
<td>[0.325]</td>
<td>[0.260]</td>
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<tr>
<td>Born 80-84</td>
<td>1.733**</td>
<td>7.287**</td>
<td>1.949**</td>
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<tr>
<td></td>
<td>[0.229]</td>
<td>[4.847]</td>
<td>[0.452]</td>
<td>[0.443]</td>
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<tr>
<td>Born 85-95</td>
<td>2.178**</td>
<td>8.760**</td>
<td>2.847**</td>
<td>1.436</td>
</tr>
<tr>
<td></td>
<td>[0.328]</td>
<td>[5.805]</td>
<td>[0.721]</td>
<td>[0.582]</td>
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<tr>
<td><strong>Race (ref: White)</strong></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Hispanic</td>
<td>0.692**</td>
<td>2.168+</td>
<td>0.738+</td>
<td>1.205</td>
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<td>[0.078]</td>
<td>[0.862]</td>
<td>[0.115]</td>
<td>[0.366]</td>
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<tr>
<td>Black</td>
<td>0.860+</td>
<td>1.380</td>
<td>0.800+</td>
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</tr>
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<td>[0.077]</td>
<td>[0.467]</td>
<td>[0.108]</td>
<td>[0.296]</td>
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<tr>
<td>Other</td>
<td>0.609**</td>
<td>1.666</td>
<td>1.117</td>
<td>1.728+</td>
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<tr>
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<td>[0.110]</td>
<td>[0.821]</td>
<td>[0.275]</td>
<td>[0.560]</td>
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<tr>
<td><strong>Immigrant</strong></td>
<td>0.451**</td>
<td>1.533</td>
<td>0.824</td>
<td>0.978</td>
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<td>[0.085]</td>
<td>[0.777]</td>
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<td>[0.328]</td>
</tr>
<tr>
<td><strong>Hisp x Immigrant</strong></td>
<td>0.376**</td>
<td>0.134*</td>
<td>0.349**</td>
<td>0.304*</td>
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<td><strong>Mother's Educ. (ref: &lt; HS)</strong></td>
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+ p<.10   * p<.05   ** p<.01
Standard errors in brackets.
Indicator variables for ages 18-45 were also included in all models, but
Odds Ratios not shown.
<table>
<thead>
<tr>
<th></th>
<th>Cohort 66-74</th>
<th>Cohort 75-79</th>
<th>Cohort 80-84</th>
<th>Cohort 85-95</th>
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<tr>
<td><strong>Sexual Behavior</strong></td>
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</tr>
<tr>
<td>Same- and other-sex partners ever</td>
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<td>0.143</td>
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<tr>
<td>Only same-sex partners ever</td>
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<td>0.007</td>
<td>0.012</td>
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<tr>
<td>Same- and other-sex partners in past year ¹</td>
<td>0.028</td>
<td>0.030</td>
<td>0.037</td>
<td>0.038</td>
</tr>
<tr>
<td>Only same-sex partners in past year ¹</td>
<td>0.009</td>
<td>0.014</td>
<td>0.019</td>
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<td><strong>Sexual Orientation (Identity)</strong></td>
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<tr>
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<td>Gay</td>
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<td>0.011</td>
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<td><strong>Sexual Attraction</strong></td>
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<tr>
<td>Bisexual ¹</td>
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<td>0.160</td>
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<tr>
<td>Gay ¹</td>
<td>0.004</td>
<td>0.006</td>
<td>0.011</td>
<td>0.011⁺</td>
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</table>

Note: Predicted probabilities are calculated from logistic regression models in Table 2, or not shown, using an average marginal effects approach to adjust for covariates.

¹ These predicted probabilities are from logistic regression models not shown, the same as those in Table 2, but changing the dependent variable.

* The last cohort is significantly higher than the first; p<.05, 2-tailed test.
+ The last cohort is significantly higher than the first; .10> p≥.05.
Table 4: Odds Ratios from Four Logistic Regressions Predicting Whether Men:
1) Have Had Male and Female Sexual Partners (Ever),
2) Have Had Only Male Sexual Partners (Ever),
3) Identify as Bisexual, and 4) Identify as Gay

<table>
<thead>
<tr>
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<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Cohort (ref: 66-74)</strong></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Born 75-79</td>
<td>0.711+</td>
<td>0.683</td>
<td>0.900</td>
<td>0.508*</td>
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<td>[0.143]</td>
<td>[0.262]</td>
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<td>Born 80-84</td>
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<tr>
<td>Born 85-95</td>
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<td>[0.271]</td>
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<td><strong>Race (ref: White)</strong></td>
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<tr>
<td>Hispanic</td>
<td>1.103</td>
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<tr>
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<td>[0.229]</td>
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<td><strong>Hisp x Immigrant</strong></td>
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<tr>
<td>HS/Some College</td>
<td>1.095</td>
<td>2.562**</td>
<td>0.686</td>
<td>1.541+</td>
</tr>
<tr>
<td></td>
<td>[0.188]</td>
<td>[0.867]</td>
<td>[0.199]</td>
<td>[0.342]</td>
</tr>
<tr>
<td>BA or more</td>
<td>1.092</td>
<td>1.780</td>
<td>0.842</td>
<td>1.092</td>
</tr>
<tr>
<td></td>
<td>[0.258]</td>
<td>[0.677]</td>
<td>[0.283]</td>
<td>[0.304]</td>
</tr>
</tbody>
</table>

N 15242 15238 15201 15201

+ p<.10  * p<0.05  ** p<0.01
Standard errors in brackets.
Indicator variables for ages 18-45 were also included in all models, but Odds Ratios not shown.
<table>
<thead>
<tr>
<th>Sexual Behavior</th>
<th>Cohort 66-74</th>
<th>Cohort 75-79</th>
<th>Cohort 80-84</th>
<th>Cohort 85-95</th>
</tr>
</thead>
<tbody>
<tr>
<td>Same- and other-sex partners ever</td>
<td>0.048</td>
<td>0.035</td>
<td>0.046</td>
<td>0.043</td>
</tr>
<tr>
<td>Only same-sex partners ever</td>
<td>0.011</td>
<td>0.008</td>
<td>0.010</td>
<td>0.015</td>
</tr>
<tr>
<td>Same- and other-sex partners in past year¹</td>
<td>0.007</td>
<td>0.009</td>
<td>0.010</td>
<td>0.014</td>
</tr>
<tr>
<td>Only same-sex partners in past year¹</td>
<td>0.020</td>
<td>0.011</td>
<td>0.016</td>
<td>0.024</td>
</tr>
<tr>
<td>Sexual Orientation (Identity)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Bisexual</td>
<td>0.013</td>
<td>0.012</td>
<td>0.021</td>
<td>0.024</td>
</tr>
<tr>
<td>Gay</td>
<td>0.028</td>
<td>0.014</td>
<td>0.016</td>
<td>0.017</td>
</tr>
<tr>
<td>Sexual Attraction</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Bisexual¹</td>
<td>0.061</td>
<td>0.045</td>
<td>0.053</td>
<td>0.058</td>
</tr>
<tr>
<td>Gay¹</td>
<td>0.019</td>
<td>0.007</td>
<td>0.009</td>
<td>0.018</td>
</tr>
</tbody>
</table>

Note: Predicted probabilities are calculated from logistic regression models in Table 4, or not shown, using an average marginal effects approach to adjust for covariates. None of the predicted probabilities for the last cohort is significantly higher than for the first cohort.

¹ These predicted probabilities are from logistic regression models not shown, the same as those in Table 4, but changing the dependent variable.