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Convergence over time or not? U.S. wages by sexual orientation, 2000–2019 ☆

Christopher Jepsen [a,*], Lisa Jepsen [b]

[a] School of Economics, University College Dublin, Ireland
[b] Department of Economics, College of Business, University of Northern Iowa, 1227 W. 27th, Cedar Falls, Iowa, 50614, United States

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ABSTRACT

An extensive literature on labor-market outcomes by sexual orientation finds lower wages for men in same-sex couples and higher wages for women in same-sex couples compared to their counterparts in different-sex couples. Previous studies analyzing multiple time periods provide suggestive evidence that the wage penalty for men in same-sex couples is heading toward zero. Using data from the American Community Survey on individuals in couples from 2000 to 2019, we find no evidence that wages, earnings, or incomes of men in same-sex couples are improving relative to married men in different-sex couples. For women in same-sex couples, we see mixed evidence of convergence relative to married women in different-sex couples. The persistence of a wage penalty for men in same-sex couples is concerning in the face of anti-discrimination policies and rising overall tolerance by Americans with respect to sexual orientation.

1. Introduction

Starting with Badgett (1995), a sizable literature examines differences in labor-market outcomes by sexual orientation. Until recently, the results from this literature have been quite consistent: lesbians earn more than heterosexual women, and gay men earn less than heterosexual men, all else equal (Klawitter, 2015; Valfort (2017); and others).[1]

Yet society has changed dramatically since the 1990s. Same-sex marriage is legal in the United States and 29 other countries (Human Rights Campaign, 2021).[2] Public opinion polls show more tolerant attitudes toward same-sex marriage and sexual orientation more generally. For example, Hansen et al. (2020b) document an increase in tolerance from 38% to 56% between 2003 and 2015 using General Social Survey data. Leo Varadkar, the Taoiseach, or prime minister, of Ireland from 2017 to 2020, is openly gay, as was one of the 2020 candidates for the President of the U.S. Given this increase in acceptance of same-sex couples, along with the lack of profitability in general of discriminatory practices, convergence in economics by sexual orientation seems likely if not inevitable.

Given the societal changes, including the increasing market pressures and legal protections against discrimination, is the gap in wages and earnings by sexual orientation diminishing? Klawitter's (2015) meta-analysis finds that the gap is smaller in recent studies compared to earlier studies. Badgett et al. (2021) find lower hourly wages for gay men in American Community Survey (ACS) data, with no evidence of any convergence in wages. In contrast, Carpenter and Eppink (2017) find an earnings premium for gay men in the 2013 to 2015 National Health Interview Survey (NHIS), and they find an earnings premium for lesbians during the same time period. Using data in the National Health and Nutrition Examination Survey from 1988 to 2007 on men who live alone, Clarke and Sevak (2013) find a shrinking gap in earnings for gay men relative to heterosexual men, although the difference is not always statistically significant. Similarly, the earnings disadvantage for gay men is diminishing in Canada (Dilmaghani 2017; Mueller 2014). There is no evidence of a gay earnings disadvantage in the UK in recent work (Aksoy et al., 2018).

Abbreviation: ACS, American Community Survey.

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* Corresponding author at: School of Economics, University College Dublin, Ireland.
E-mail address: christopher.jepsen@ucd.ie (C. Jepsen).

[1] Klawitter (2015) and Valfort (2017) summarize the literature on earnings by sexual orientation.

[2] As of October 2021, the countries are: Argentina, Australia, Austria, Belgium, Brazil, Canada, Colombia, Costa Rica, Denmark, Ecuador, Finland, France, Germany, Iceland, Ireland, Luxembourg, Malta, Mexico, the Netherlands, New Zealand, Norway, Portugal, South Africa, Spain, Sweden, Switzerland, Taiwan, the United Kingdom, the United States of America, and Uruguay.

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For women, Badgett et al. (2021) find a declining hourly wage premium for lesbians in the ACS for 2000 to 2018, and Cushing-Daniels and Yeung (2009) find a declining lesbian earnings premium in the General Social Survey between 1988 and 2006, which is consistent with the UK evidence in Aksoy et al. (2018) of no earnings differences between lesbians and heterosexual women. Carpenter and Eppink (2017) document an earnings premium for lesbians in the 2013 to 2015 NHIS.

We contribute to this literature by taking a deep dive into earnings and employment by sexual orientation as inferred from relationship status (described below). First, we calculate the wage and earnings gap over a 20-year time period from 2000 to 2019, complementing recent research looking at 1988 to 2007 (Clarke and Sevak, 2013) and 2013 to 2015 (Carpenter and Eppink, 2017). Using the U.S. ACS, we have a large sample of men and women in same-sex couples, with 2400 or more full-time workers of each sexual orientation each year starting in 2005. The ACS is not without its limitations, as discussed below. The most notable is the restriction of the sample to cohabiting individuals. Consequently, like previous work looking at sexual orientation and labor-market outcomes, the analysis is descriptive and probably not causal given issues such as selection into cohabitation, employment, full-time work, and "coming out."

Second, we create four measures of earnings, including annual wages / salaries; earnings, defined as wages / salaries plus self-employment income; annual income, defined as earnings plus unearned income such as dividends; and hourly wages. By analyzing different measures of labor-market outcomes, we study the sensitivity of the results to variations in the measure of labor-market outcomes. Here, we complement the analysis of hourly wages for ACS data from 2000 to 2018 in Badgett et al. (2021).

Third, we study the sensitivity of the results to various age ranges. We expect labor-market behaviors to be correlated with age, so we control for different times in a person's earnings' life.

We find no evidence of convergence for men and mixed evidence for women since the Global Financial Crisis (GFC). For women, the raw gaps in all four measures of earnings decline dramatically between 2000 and 2019, whereas the regression-adjusted gaps have much slower convergence. By 2019, for full-time workers, men in same-sex couples have lower wages of approximately 10% compared to married men in different-sex couples, whereas women in same-sex couples have higher wages of approximately eight percent compared to married women in different-sex couples. The gap is similar for earnings and income, but it is somewhat lower for hourly wages compared to annual wages. The gap is robust to several factors such as age ranges, sample weights, and variation in control variables.

2. Data

Data are from the American Community Survey (ACS) Public Use Microdata Sample (PUMS). The ACS is the largest individual-level data set collected annually by the U.S. Census Bureau. We use data from the 2000 to the 2019 surveys. These data are available via Mendeley Data (http://dx.doi.org/10.17632/b8wbzgtrtp.1).

Sexual orientation is identified through the ACS question on relationship to head of household. Thus, we can only identify the sexual orientation of individuals who are cohabiting, either as the head of household or as the cohabiting partner / spouse. Until 2013, all same-sex couples were identified as unmarried even if they listed their marital status as married.[3] To provide consistent analysis across years, therefore, we do not distinguish between unmarried and married same-sex couples. Our sample of cohabiting individuals is divided into four mutually exclusive and exhaustive couple types: men in same-sex couples, women in same-sex couples, unmarried different-sex couples, and married different-sex couples.

The sample is restricted to provide consistent comparisons to previous research. Because the focus is on labor-market outcomes, our preferred sample includes individuals between the ages of 18 and 64 who have positive – and non-allocated – wage / salary income. We exclude individuals where either member of the couple (head or partner / spouse) has missing or allocated values for sex, relationship to head of household, or marital status. The restriction reduces the likelihood that we misclassify a different-sex couple as a same-sex couple (Black et al., 2006; Gates and Steinberger, 2009). Because our focus is on full-time workers, unless otherwise specified, the analysis is restricted to individuals who worked at least 35 h per week and worked at least 27 weeks in the previous year.

The 2000 to 2004 waves of ACS are pilot waves, with under 400,000 observations in 2000 and approximately 1.1 to 1.2 million observations per year from 2001 to 2004. In 2005, the number of people surveyed increased to nearly 2.9 million individuals. Since then, the sample size has gradually increased; in 2019, there were roughly 3.2 million individuals.

Appendix Table 1 shows the sample size of full-time workers between ages 18 and 64 (with no allocated / missing values as mentioned above) in each of the four couple types. Starting in 2013, the number of full-time workers in same-sex couples began to grow substantially, in part because the Census Bureau stopped allocating marital status of same-sex married couples in that year. In 2019, the dataset contains 5260 women in same-sex couples and 5288 men in same-sex couples, more than double the numbers in 2012. The number of married individuals in different-sex couples dropped slightly in 2008 and continued to drop until 2015. Modeling the determinants of the decision to cohabitate is beyond the scope of this paper, but the trends in our sample sizes are consistent with current trends of increasing cohabitation in lieu of legal marriage for younger generations such as Millennials (Manning, 2020; Manning et al., 2019).

Also notable are two changes in the ACS between 2007 and 2008. The first is a formatting change in the questionnaire making it more difficult for participants to mark both male and female genders accidently. The second is a set of "technological improvements in data collection by interviewers and efforts to make the processing and editing more consistent between data in the ACS and the 2010 Census" (U.S. Census Bureau, 2013, page 2).

3. Descriptive statistics for full-time workers

Before exploring differences in labor-market outcomes by couple type, we first document trends in demographic characteristics for full-time workers over the time period 2000 to 2019. Badgett et al. (2021) provide extensive information on the lesbian, gay, bisexual, transgender, and queer (LGBTQ) populations.[4]

Fig. 1 shows the trends in age by couple type. The top panel is for women, and the bottom panel is for men. For all years, the ranking of average age is similar by gender: married women / men in different-sex couples are slightly older than women / men in same-sex couples, and unmarried women / men in different-sex couples are noticeably younger. There is a general trend of increasing age over the time period for individuals in different-sex couples. For individuals in same-sex couples, average age peaks in 2013 (around 43–44 years of age) and declines by 1 to 2 years between 2014 and 2019. Note that the trend is noisier for men and women in same-sex couples, likely due to the much smaller sample sizes.

Next, we explore patterns in self-reported race and ethnicity. The race and ethnicity category in the ACS is a 'check all that apply' out-

[3] The 2012 data contain a variable to identify same-sex married couples, even though all same-sex couples are treated as unmarried.

[4] For information on the demographic trends for all cohabiting individuals, not just full-time workers, see Jepsen and Jepsen (2020).

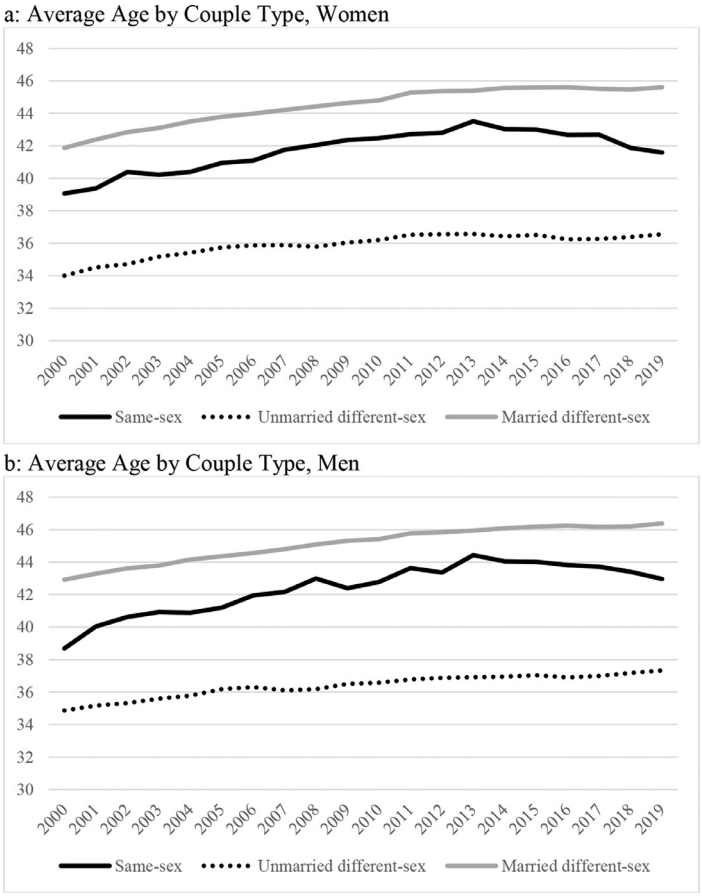


Fig. 1. Notes: Sample is for women / men working full-time (35+ hours per week, 27+ weeks per year), ages 18 to 64; no allocated values (among either partner) for sex, relationship to head of household, or marital status.

come so that individuals may report more than one race or ethnicity. Fig. 2 contains trends in the percent of individuals identifying as white.[5] Same-sex couples have a higher percentage of whites than different-sex couples. The differences between couple types have narrowed, however, from over six percentage points in 2000 to approximately two percentage points in 2019. The narrowing is primarily due to a decline in percentage white among same-sex couples, as the percentage for married, different-sex couples is slightly increasing. For all couple types, over 80% of individuals self-identify as white. Appendix Fig. 1a and 1b illustrate the trends in the percentage black, Hispanic, and other race separately by couple type for women and men, respectively. In general, the largest increase over the time period is for the percentage Hispanic, with smaller increases, if not decreases, for the percentage black.

An important determinant of wages and earnings is education. Fig. 3 illustrates the percentage of individuals with a bachelor's degree (or more) between 2000 and 2019. Between 2001 and 2009, the percentage of college graduates is between 53 and 55 for women in same-sex couples, whereas the percentage increases steadily for other couple types. Individuals in same-sex couples have the highest levels of education, between 50 and 60%, although the percentage of married women in different-sex couples with at least a bachelor's degree is only two percentage points lower than that for women in same-sex couples in 2019. Roughly one-third of married individuals in different-sex couples have bachelor's degrees in 2001, increasing to 53% for women and 45% for men by 2019. Unmarried individuals in different-sex couples have the lowest percentages, with values of 20 to 25% in 2000. By 2019, the percentages have increased to 44% for women and 31% for men.

[5] Like many U.S. studies, the ACS reports Hispanic ethnicity separately from the race categories, in that individuals are simply asked whether or not they identify as Hispanic. Thus, each Hispanic person will also be represented by at least one of three racial categories of black, white, and other.

4. Methods

The results in the previous section suggest that, in general, the trends in demographics are often similar across couple types and over time, but some clear differences between couple types exist. In this section, we present the econometric specification for estimating labor-market outcomes using ordinary least squares (OLS) regression, although – as mentioned earlier – this analysis is descriptive rather than causal. We estimate separate models for men and for women[6], and we estimate separate models for each year.[7] Eq. (1) contains the main regression specification for a given year:

$$Ln(Y_i) = \alpha + \beta * Coupletype_i + \gamma * X_i + \varepsilon_i \quad (1)$$

In the preferred specification, Y_i is annual wages for individual i, measured in natural logs. For robustness, we also estimate alternate models where annual earnings, annual income,[8] and hourly wage[9] are the dependent variables, again estimated in natural logs.

For couple type, we include two dummy variables, one for being in an unmarried, different-sex cohabiting couple, and one for being in a same-sex couple. Because the omitted couple type is different-sex married couples, the coefficients for couple type are interpreted relative to married individuals in different-sex couples. As mentioned previously,

[6] The ACS data have two categories for gender: male and female.

[7] The models are estimated separately by year to provide a more flexible model and to reduce computational burden.

[8] As mentioned previously, annual earnings are equal to the sum of (1) wage and salary earnings and (2) self-employment earnings. Income includes earnings and unearned income such as dividend payments.

[9] Hourly wage equals the annual wage divided by the number of hours in the year. In years where the weeks worked per year is categorical, we define the weeks worked as the midpoint of the category. Given this limitation on the weeks worked per year category, the preferred wage measure is annual wages.

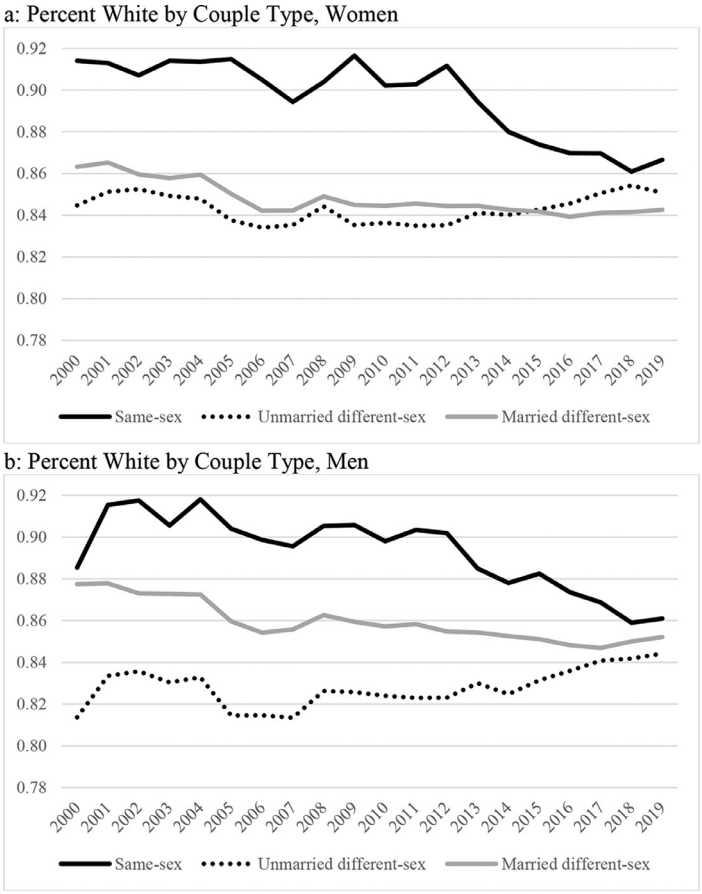


Fig. 2. Notes: Sample is for women / men working full-time (35+ hours per week, 27+ weeks per year), ages 18 to 64; no allocated values (among either partner) for sex, relationship to head of household, or marital status.

for most years the ACS does not distinguish between married and unmarried same-sex couples. For consistency across years, our preferred method combines individuals in married and unmarried same-sex couples into one overall same-sex couple category.

For each individual, the vector X contains the following observable determinants of log wages: race / ethnicity relative to the omitted category of white; age and age squared; education relative to the omitted category of high school graduates; number of children in the household, with separate variables for school-age children (ages 6 to 17) and for younger children (ages 0 to 5); occupation relative to the omitted category of managerial occupations; industry relative to the omitted category of education, medical, family services, and administration; a dummy variable for disabled; a dummy variable for currently in school; and state fixed effects. We assume that heteroskedasticity exists and use Stata's "robust" command to adjust the standard errors accordingly.

Finally, we also estimate two models of labor supply. In the first, the dependent variable is the usual number of hours worked per week.[10] The regression sample for this model is all workers (full and part time) in order to isolate the intensive margin of hours worked among workers from the extensive margin of employment. Given the limited variation in hours worked among full-time workers, the regression sample for this model is all workers – individuals with positive, non-allocated values for hours worked.

Second, we estimate a linear probability model on employment. The dependent variable is equal to one for individuals who are employed and zero for individuals who are unemployed or who are not in the labor force. Employment is determined by the ACS created variable called employment status recode.

In all models, allocated or imputed values for the dependent variable are treated as missing. Bollinger and Hirsch (2013) document concerns with using imputed observations in the Current Population Survey (CPS), and the ACS uses the same imputation procedure as the CPS. To summarize, the preferred annual log wages analysis is limited to full-time workers with positive, non-allocated values of annual wages. Similarly, the analyses for earnings and income are limited to full-time workers with positive, non-allocated values of annual earnings and income, respectively. The hourly wage analysis is limited to full-time workers with positive hourly wages and non-allocated values of annual wages, hours worked per week, and weeks worked per year. When the dependent variable is hours worked per week, the regression sample is limited to people with positive, non-allocated values of hours worked (e.g. both full-time and part-time workers). Finally, the employment analysis is limited to people who have non-allocated values of the ACS-generated employment recode variable; this sample includes workers, unemployed individuals, and individuals not in the labor market.

5. Results

5.1. Raw differences

Fig. 4 contains the raw differences in labor-market outcomes by couple type. Specifically, the figure contains the coefficients from regressions (as in Eq. (1)) that include couple type variables but exclude demographic variables (vector X). The top panel is for women, and each line represents the coefficient for women in same-sex couples compared to the omitted group of married women in different-sex couples. The bot-

[10] We also estimate a model where hours worked per year is the dependent variable, again for workers. However, in most years, the weeks worked per year is a categorical variable (with 6 categories). Consequently, the measure of hours worked per year is less precisely estimated that the usual hours worked per week.

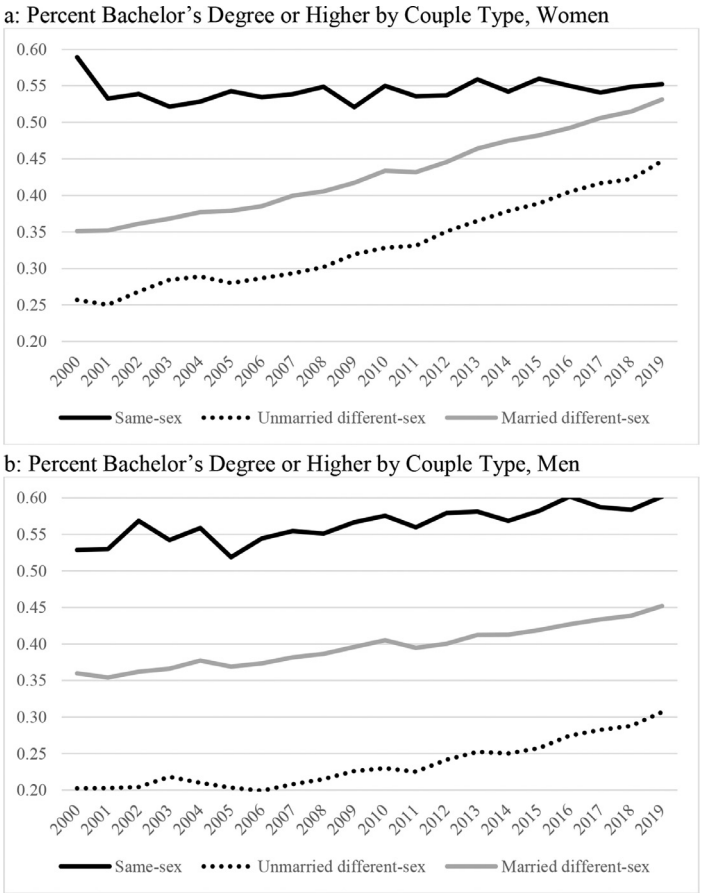


Fig. 3. Notes: Sample is for women / men working full-time (35+ hours per week, 27+ weeks per year), ages 18 to 64; no allocated values (among either partner) for sex, relationship to head of household, or marital status.

tom panel is for men, and the reported coefficient is for men in same-sex couples compared to the omitted group of married men in different-sex couples. Recall that the results for all the wage, earnings, or income regressions are for the sample of full-time workers, defined as working at least 35 h a week for at least 27 weeks (e.g. more than half the year) in the last 12 months.

The raw gap between women in same-sex couples and married women in different-sex couples (the reference group) declines dramatically over the twenty-year period. In 2000, women in same-sex couples have better labor-market outcomes ranging from 0.22 log points for hourly wages to 0.37 log points for income. By 2019, the raw advantage is less than 0.05 log points for three of the four outcomes. Although mean outcomes improve for both couple types, the increase in mean outcomes is steeper for married women in different-sex couples. This convergence in labor-market outcomes mirrors a convergence in race / ethnicity (Fig. 2) and education levels (Fig. 3), suggesting a potential role for these characteristics in explaining the convergence in outcomes. In contrast, for men, the raw gap in labor-market outcomes between men in same-sex couples and men in different-sex couples is indistinguishable from zero for most years.

For women, Table 1 contains the regression coefficients and standard errors for couple type for the regressions – including demographic characteristics (vector X) – depicted in Eq. (1). The dependent variable is log annual wages. For women in same-sex couples, the wage advantage is between 0.075 log points (in 2019) and 0.120 log points (in 2001), a significant decline under the standard assumptions for comparing coefficients across different samples. Among women in different-sex couples, unmarried women have lower wages of 1–3 log points (with the exception of 2000 and 2001, where wages are statistically indistinguishable).

The coefficients and standard errors for men are reported in Table 2. Men in same-sex couples have lower wages, ranging from 0.081 log points in 2014 to 0.160 log points in 2003. The gap in wages is as large, if not larger in some years, as that between married and unmarried men in different-sex couples.

5.2. *Preferred specification*

Fig. 5 presents the coefficients for same-sex couples in regressions of four different labor-market outcomes: log annual wages, log annual earnings, log annual income, and log hourly wage. Again, the comparison group consists of married individuals in different-sex couples. Consistent with much of the previous literature, women in same-sex couples have a wage premium, and men in same-sex couples have a wage penalty. However, the size of the premium / penalty varies by outcome, year, and sex.

Between 2000 and 2008, the pattern is roughly similar between men and women. However, as noted by the U.S. Census Bureau (2013), some of the same-sex couples may be misclassified prior to 2008, so the results from these years should be interpreted with caution. The coefficients fluctuate between 2000 and 2004, during the pilot studies of the ACS with relatively small samples of full-time workers in same-sex couples (Appendix Table 1). Between 2005 and 2008, the premium for women in same-sex couples rises (except for income), whereas the size of the penalty declines for men in same-sex couples.

Between 2008 and 2019, the premium for women in same-sex couples declines for most years, with a cumulative decline of 0.03 to 0.04 log points. For example, the income premium declines from 0.119 to 0.104, whereas the earnings premium declines from 0.108 to 0.075. For men, we see a different pattern. The penalty for same-sex couples con-

Table 1
Female couple type coefficients for log annual wages, 2000–2019 ACS.

	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Same-sex couple	0.109	0.120	0.084	0.113	0.083	0.086	0.105	0.117	0.101	0.112
	(0.029)	(0.017)	(0.019)	(0.019)	(0.017)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Different-sex unmarried	-0.012	-0.001	-0.018	-0.020	-0.012	-0.018	-0.020	-0.024	-0.025	-0.022
	(0.011)	(0.006)	(0.006)	(0.006)	(0.006)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Observations	30,648	99,280	90,495	98,403	100,462	239,584	242,344	244,873	250,102	246,205
R-Squared	0.319	0.344	0.341	0.341	0.335	0.341	0.346	0.352	0.355	0.359
	2010	2011	2012	2013	2014	2015	2016	2017	2018	2019
Same-sex couple	0.098	0.089	0.086	0.093	0.107	0.100	0.096	0.087	0.084	0.075
	(0.011)	(0.012)	(0.011)	(0.010)	(0.010)	(0.009)	(0.008)	(0.009)	(0.008)	(0.008)
Different-sex unmarried	-0.027	-0.019	-0.030	-0.025	-0.025	-0.029	-0.025	-0.021	-0.020	-0.024
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Observations	233,307	224,058	227,796	228,177	226,791	231,977	235,272	241,636	244,616	245,737
R-Squared	0.361	0.365	0.365	0.356	0.359	0.358	0.359	0.353	0.353	0.354

Notes: Separate regressions are estimated for each year. In addition to the coefficients listed, all models contain additional control variables as outlined in the methods section. The coefficients for different-sex unmarried in 2000 and 2001 are not statistically significantly different from zero at the one-percent level for a two-sided test, whereas the coefficient for 2004 is statistically significantly different from zero at the five-percent level for a two-sided test. All other coefficients are statistically significantly different from zero at the one percent level for a two-sided test. Standard errors are robust to heteroskedasticity.

Table 2
Male couple type coefficients for log annual wages, 2000–2019 ACS.

	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Same-sex couple	-0.143	-0.106	-0.097	-0.160	-0.148	-0.120	-0.112	-0.101	-0.114	-0.117
	(0.031)	(0.020)	(0.022)	(0.019)	(0.019)	(0.012)	(0.012)	(0.013)	(0.013)	(0.013)
Different-sex unmarried	-0.144	-0.140	-0.135	-0.150	-0.138	-0.139	-0.138	-0.134	-0.141	-0.149
	(0.011)	(0.006)	(0.006)	(0.006)	(0.006)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Observations	45,246	146,418	132,977	144,501	147,236	352,797	348,964	355,301	356,776	344,877
R-Squared	0.315	0.321	0.321	0.312	0.317	0.325	0.315	0.334	0.334	0.336
	2010	2011	2012	2013	2014	2015	2016	2017	2018	2019
Same-sex couple	-0.085	-0.089	-0.089	-0.088	-0.081	-0.121	-0.102	-0.101	-0.113	-0.110
	(0.013)	(0.013)	(0.013)	(0.010)	(0.011)	(0.010)	(0.010)	(0.010)	(0.009)	(0.009)
Different-sex unmarried	-0.142	-0.142	-0.140	-0.139	-0.135	-0.135	-0.134	-0.132	-0.132	-0.128
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Observations	321,107	309,897	317,177	316,800	315,060	319,633	322,290	328,871	331,017	328,842
R-Squared	0.346	0.352	0.348	0.343	0.342	0.342	0.341	0.336	0.334	0.328

Notes: Separate regressions are estimated for each year. In addition to the coefficients listed, all models contain additional control variables as outlined in the methods section. All coefficients are statistically significantly different from zero at the one-percent level for a two-sided test. Standard errors are robust to heteroskedasticity.

verges toward zero between 2008 (if not earlier) and 2014, as well as between 2015 and 2017. But the penalty diverges from zero between 2014 and 2015 and between 2017 and 2019. By 2019, the penalty is roughly equal to the penalty in 2008.

Across outcomes, a consistent pattern occurs for men and women. The smallest penalties and premia are for hourly wages. Prior to 2008, the largest penalties and premia are for income.[11] Starting in 2008, the penalties and premia are quite similar between annual wages, annual earnings, and annual income.

Appendix Fig. 2 contains regression results for the full sample of workers rather than the subsample of full-time workers. The wage premia for women in same-sex couples and the penalties for men in same-sex couples are still present in the full sample and follow the same temporal patterns. For the three annual measures, the premia and penalties are noticeably larger in the full sample. The hourly wage results are quite similar between the full-time sample (Fig. 5) and the sample of all workers (Appendix Fig. 2), consistent with a similar hourly wage penalty / premium for full-time workers and part-time workers.

5.3. Labor supply

The lower annual wage premia and penalties for full-time workers relative to all workers suggests the possibility of labor-supply differences by sexual orientation. A substantial literature starting with Tebaldi and Elmslie (2006) documents a labor-supply penalty for men in same-sex couples and a premium for women in same-sex couples, the same pattern we observe for wages.[12] Fig. 6 contains the results for two measures of labor supply. The first measure is the usual hours worked per week, and the regression sample is the set of all workers regardless of full-time or part-time status. The second measure is a dummy variable for employment, and the regression sample is all individuals,[13] including individuals who are unemployed or not in the labor force. The scale on the left is for hours worked, and the scale on the right is for employment probability.

Women in same-sex couples work more hours per week and have higher employment probabilities than married women in different-sex couples. The premium for hours worked declines from its largest gap of 3.5 h per week in 2003 to 2.5 h per week by 2019. In contrast, the employment premium is around 0.08 for most of the time period and does

[11] When we use a common sample for all wage / income outcomes (results available from the authors upon request), the coefficients for income are nearly identical to those for earnings or wages. This common sample has no allocated values for wages, earnings, income, or hours worked. Thus, the coefficients for income may be driven by individuals who have allocated values of wages, earnings, or hours but not income.

[12] Recent work concerning sexual orientation and labor supply looks closely at the role of factors such as tolerance (Hansen et al., 2020b) and state and local policies (Delhommer, 2020).

[13] As mentioned above, we exclude individuals with allocated values for employment.

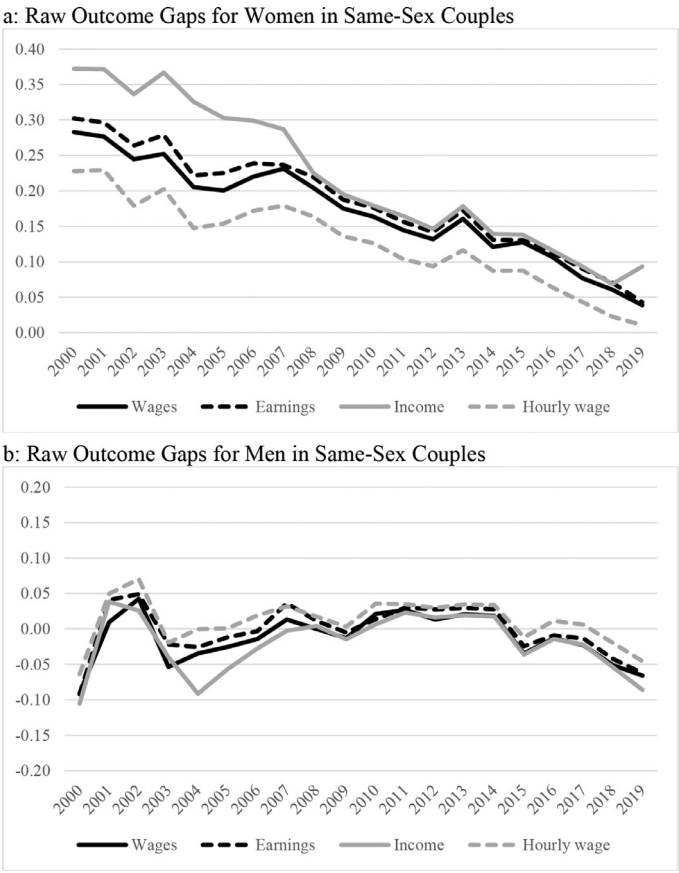


Fig. 4. Notes: Each point is the coefficient for women / men in same-sex couples (compared to married women / men in different-sex couples) from a separate regression where the only coefficients are dummy variables for couple type. The label is the dependent variable, measured in natural logs.

not demonstrate any convergence toward zero. Thus, any convergence in labor supply for women in same-sex couples is occurring at the intensive margin of hours worked for individuals already working rather than at the extensive margin between being employed and not being employed.

Men in same-sex couples have fewer hours worked and a lower likelihood of employment compared to married men in different-sex couples. In 2019, the gap between these two groups is 1.76 h per week and five percentage points for employment. For comparison, Carpenter and Eppink (2017) find a similar result of four to five percentage points in NHIS data. In Fig. 6, few trends across the twenty-year period emerge aside from a decline in both gaps between 2017 and 2019.

5.4. Time trend

To investigate the possibility of a time trend in the wage premium and penalty, we run simple regressions of a yearly time trend on the regression coefficients in Fig. 5. The regressions are limited to the years between 2008 and 2019 due to concerns about data quality for same-sex couples prior to 2008.[14] Eq. (2) illustrates the regression specification:

$$\hat{\beta} = \theta + \lambda \cdot YEAR + \nu \quad (2)$$

[14] We run this simple regression rather than a pooled, individual-level regression because the individual-level regression would contain millions of observations, and ALL the regression coefficients would be statistically significant from zero. In other words, we would find a statistically significant time trend but perhaps not an economically significant one.

Table 3
Coefficient on time trend for same-sex couple coefficients, 2008–2019.

	Annual Wages	Annual Earnings	Annual Income	Hourly Wages
Women				
Coefficient	-0.0018	-0.0021	-0.0020	-0.0022
Standard Error	0.0007	0.0007	0.0007	0.0007
T-stat	-2.524	-2.979	-2.763	-2.997
Observations	12	12	12	12
R-squared	0.389	0.471	0.432	0.473
Years until Convergence	42	35	53	25
Men				
Coefficient	-0.0006	-0.0005	-0.0022	-0.0009
Standard Error	0.0012	0.0012	0.0014	0.0011
T-stat	-0.489	-0.389	-1.502	-0.828
Observations	12	12	12	12
R-squared	0.023	0.015	0.184	0.064
Years until Convergence	N/A	N/A	N/A	N/A

Notes: Each panel contains the results from a separate regression, for a total of eight regressions. The dependent variable is the coefficient for being in a same-sex couple, as illustrated in Fig. 4 (and in Tables 1 and 2 for log annual wages).

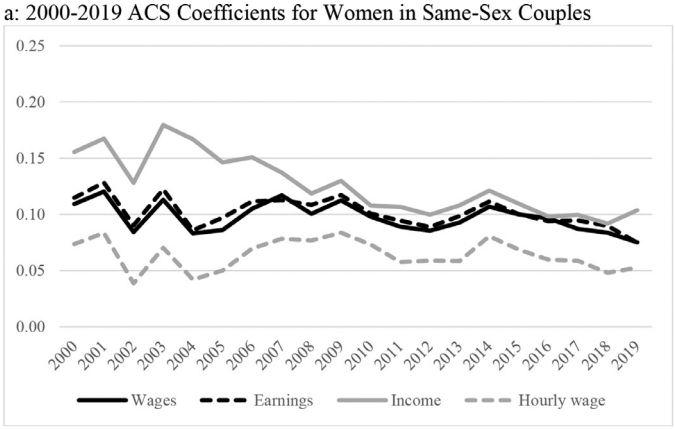
is the regression coefficient from Fig. 5 (Tables 1 and 2 when the dependent variable is log annual wages), and YEAR is the calendar year.

The results from these regressions are reported in Table 3. The results provide evidence of convergence among women, as the coefficient for a time trend is always statistically significantly different from zero at the five-percent level (for a two-sided test). The rate of convergence is approximately -0.002 log points (or -0.002%). In the final row of each panel, we estimate the number of years until convergence assuming that the premium in 2019 declines by the estimated coefficient in each year thereafter. For example, for log annual wages, if the premium of 0.075 converges by 0.0018 each year, it will converge to zero in 42 years. The years until convergence varies from 25 years to 53 years. If we include all the coefficients from 2000 to 2019, the time until convergence is even longer, ranging from 27 to 137 years. We interpret these findings as suggestive of slow *economic* convergence despite back-of-the-envelope evidence of *statistical* convergence. To put this time period in perspective, the Stonewall riots occurred in 1969, 50 years before the end data point of 2019. The rate of convergence is much slower than that found in the meta-analysis by Klawitter (2015).

Table 3 provides no evidence of convergence in the penalty for men in same-sex couples based on the regression coefficients for the years between 2008 and 2019. If we use all the coefficients from 2000 to 2019, the years until convergence are between 52 (log annual income) and 175 years (log hourly wages). The pattern for men is also at odds with the trends found in other data sets (Carpenter and Eppink, 2017; Clarke and Sevak, 2013) or in a meta-analysis (Klawitter, 2015).

Why would the results for men in same-sex couples differ for the ACS relative to other data sets? First and foremost, the ACS data only contain men and women in same-sex couples, and these individuals may differ from the broader population of gay men and lesbians. In the NHIS data used by Carpenter and Eppink (2017), 66.1% of lesbians are cohabiting, compared with 47.4% of gay men. The descriptive statistics reported in their Table 1 are similar to the demographic trends we report in Figs. 1 to 3 for age and education; the percentage white is somewhat higher in the ACS than in the NHIS. Carpenter and Eppink (2017) and Clarke and Sevak (2013) use individual-level data on sexual orientation rather than the approach in the ACS data of inferring sexual orientation from the genders of the head of household and the partner / spouse. Although Carpenter and Eppink (2017) find no evidence of a wage penalty for gay men when they limit the sample to partnered men, the number of partnered gay men in their sample is 146. The sample of gay men – partnered or not – is only 77 in Clarke and Sevak (2013). Thus, it is difficult to determine in currently-available data whether the pattern of results

a: 2000-2019 ACS Coefficients for Women in Same-Sex Couples



b: 2000-2019 ACS Coefficients for Men in Same-Sex Couples

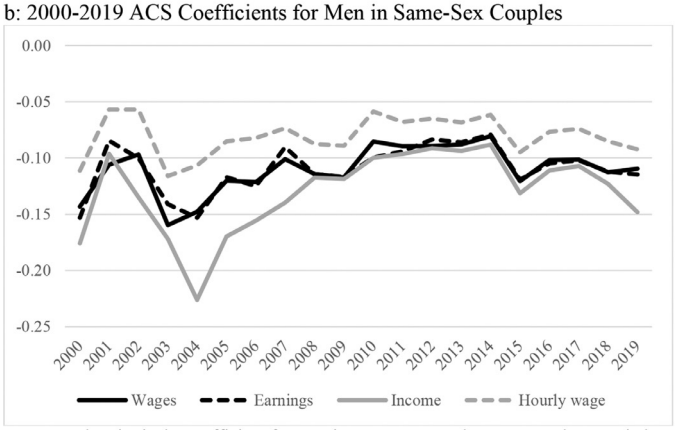
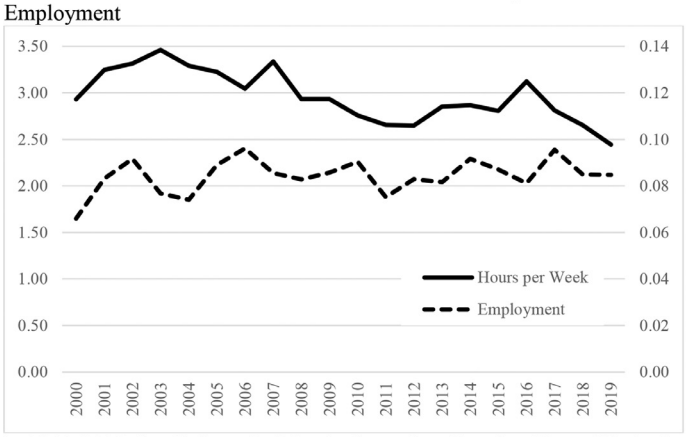


Fig. 5. Notes: Each point is the coefficient for individuals in same-sex couples (compared to married individuals in different-sex couples) from a separate regression, based on Eq. (1). The label is the dependent variable, measured in natural logs. All coefficients are statistically different from zero at the one-percent level for a two-sided test.

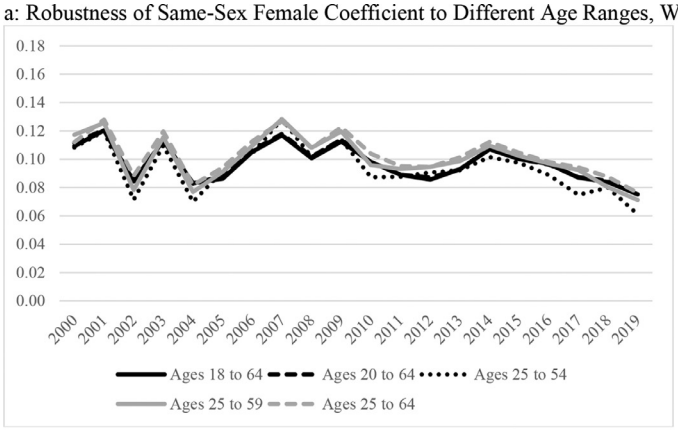
a: 2000-2019 Coefficients for Women in Same-Sex Couples, Hours Worked and Employment



b: 2000-2019 Coefficients for Men in Same-Sex Couples, Hours Worked and Employment

Fig. 6. Notes: Each point is the coefficient for individuals in same-sex couples (compared to married individuals in different-sex couples) from a separate regression, based on levels rather than log-specification of Eq. (1). The dependent variables are usual hours per week and employment; the latter is estimated as a linear probability model. All coefficients are statistically different from zero at the one-percent level for a two-sided test.

a: Robustness of Same-Sex Female Coefficient to Different Age Ranges, Wages



b: Robustness of Same-Sex Male Coefficient to Different Age Ranges, Wages

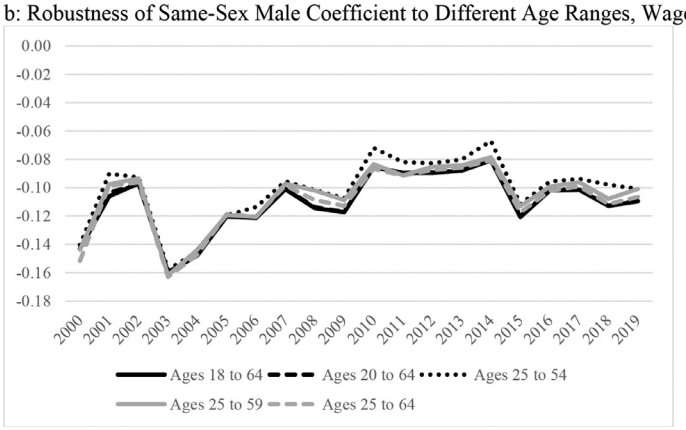


Fig. 7. Notes: Each point is the coefficient for individuals in same-sex couples (compared to married individuals in different-sex couples) from a separate regression, based on Eq. (1). The dependent variable is log annual wages. All coefficients are statistically different from zero at the one-percent level for a two-sided test.

by sexual orientation differs between individuals who are cohabiting versus those who do not live with a partner.

5.5. Robustness checks and extensions

An unlikely explanation for the difference in results is variation in the dependent variable to capture labor-market outcomes. The results in Fig. 5 illustrate a similarity across different outcome measures, including the annual earnings measure used in Carpenter and Eppink (2017). Clarke and Sevak (2013) use household earnings, as their data set does not contain individual earnings. Although the results for hourly wages suggest a smaller penalty for men in same-sex couples, the pattern of convergence – or lack thereof – is similar when compared to annual wages, earnings, or income.

The age range of the sample also differs across studies of earnings by sexual orientation. Our results are for ages 18 to 64, the same as Blandford (2003) and Carpenter (2004). Other age ranges in the literature include: 25 to 64 in Carpenter and Eppink (2017), 18 to 59 in Clarke and Sevak (2013); 25 to 59 in Antecol et al. (2008); 25 to 54 in Antecol and Steinberger (2013); 20 to 64 in Allegretto and Arthur (2001); and 18 to 65 in Jepsen (2007).[15]

Fig. 7 contains the results across age ranges for the annual wage regressions; results for annual earnings, annual income, and hourly wages follow the same pattern and are available from the authors upon request. The results are quite similar across the different age ranges for both men and women. For many years, the penalty for men in same-sex couples and the premium for women appear to be slightly smaller for prime-age earners, ages 25 to 54. Thus, the different age ranges used across papers is most likely not the explanation for differences in results.

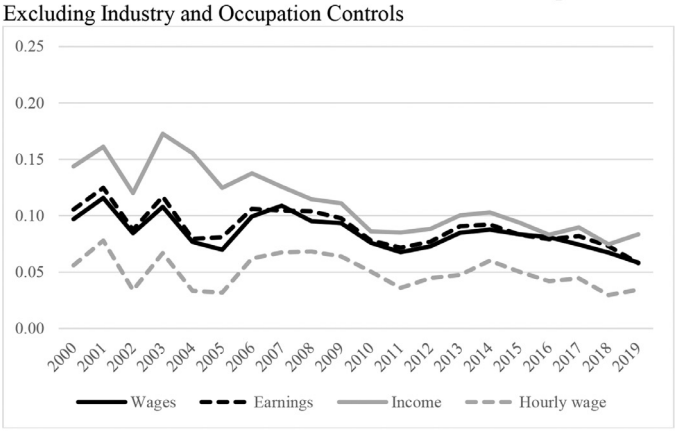
Plug et al. (2014) show that, in Australia, gay men and lesbians avoid certain occupations identified as prejudiced against sexual minorities. In the U.S., Del Río and Alonso-Villar (2019) document occupational segregation for same-sex couples. We explore the sensitivity of our results to the exclusion of control variables for industry and occupation. The results from this more parsimonious model are illustrated in Fig. 8. The premia for women in same-sex couples are somewhat lower in the models that exclude controls for occupation and industry. This result is consistent with the comparisons between women in same-sex couples and women in different-sex couples in Del Río and Alonso-Villar (2019). In contrast, Antecol et al. (2008) find little role of occupational sorting in explaining wage differentials in the 2000 Census, where they use Oaxaca-Blinder decompositions. Perhaps the role of occupation has changed over time, as the role of occupation is generally smaller at the start of our period (close to 2000) compared to later years (such as the 2010–2015 time period studied by Del Río and Alonso-Villar, 2019).[16]

For men, the results are nearly identical whether or not the model includes controls for occupation and industry. This result is consistent with Antecol et al. (2008). Although Martell (2018) finds similar results between models with and without controls for occupation, he shows that the penalty for men in same-sex couples is smaller in occupations with more independence. Our results for men differ from Del Río and Alonso-

[15] Rather than attempt to cover all the age ranges and all the papers in the vast literature on labor-market outcomes by sexual orientation, this list focuses on studies that either look at labor-market outcomes over time (Carpenter and Eppink, 2017; Clarke and Sevak, 2013) or that use Census or ACS data. Even so, the list is far from exhaustive.

[16] Delhommer (2020) finds that the penalty for men in same-sex couples and premium for women in same-sex couples are smaller in models that include occupation fixed effects. Because the paper contains interaction terms between state and local anti-discrimination laws and sexual orientation (as well as county fixed effects), we do not attempt to make a direct comparison between his results and ours.

a: 2000-2019 ACS Coefficients for Women in Same-Sex Couples Excluding Industry and Occupation Controls



b: 2000-2019 ACS Coefficients for Men in Same-Sex Couples Excluding Industry and Occupation Controls

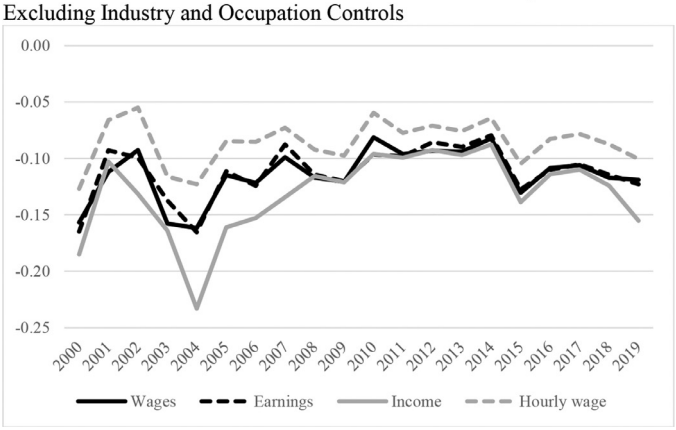


Fig. 8. Notes: Each point is the coefficient for individuals in same-sex couples (compared to married individuals in different-sex couples) from a separate regression, based on Eq. (1). The regressions exclude controls for occupation or industry, unlike the regressions for all other figures. The label is the dependent variable, always measured in natural logs. All coefficients are statistically different from zero at the one-percent level for a two-sided test.

Villar (2019)), who find that occupational sorting explains a sizeable portion of the penalty for men in same-sex couples.

Another robustness check is to see if our results are sensitive to the inclusion or exclusion of unmarried individuals in different-sex couples within the comparison group. Appendix Fig. 3 shows the results when the comparison group is all individuals in different-sex couples (married and unmarried). The coefficients for women and men in same-sex couples are nearly identical to the coefficients in Fig. 5, showing that the results are not sensitive to the inclusion or exclusion of unmarried individuals in the comparison group.

The next robustness test is to compare the results from different samples. Fig. 9 illustrates the results from four different samples. The first is the preferred sample used throughout the paper. The second sample excludes the top 1% of the dependent variable in order to limit the influence of outliers, as well as to address the top-coding of labor-market outcomes in the ACS. The third is limited to white, U.S. born individuals only, in order to isolate the effects of sexual orientation from the effects of race / ethnicity and immigration. The fourth sample is limited to individuals who live in households without any children ages 0 to 17. Because the pattern of results is similar for all wage measures, Fig. 9 focuses on the results for log annual wages.[17]

For women, the pattern of results is similar across the four samples, and, starting in 2005 when over one million people are surveyed, all the estimated premia for women in same-sex couples are within 0.02 of each other. The premium is slightly lower when the sample is limited to white, U.S.-born individuals. Such a result is inconsistent with the notion that lesbians who are people of color face a double disadvantage based on race and sexual orientation.[18] The premium is also slightly smaller when excluding the top 1%, illustrating that the premia is influenced a bit by outliers in the form of very high earners. The similarity of results between the preferred sample and the sample without children in the household demonstrates that differences by sexual orientation for households with children are not driving the premium. The controls for children in the ACS data are imprecise, however, so future work with better data on parenthood is needed.

Although the temporal pattern is similar across samples, the size of the penalty for men in same-sex couples varies with the sample used. Compared to the full sample, the penalty is larger for the sample of U.S.-born whites. As with women, this result is not consistent with a story of men of color in same-sex couples facing a double disadvantage due to race and sexual orientation. The penalty is smaller when we restrict the sample to men in households without children. Such a finding is plausible if any child penalty in different-sex couples is borne by the woman (presumably the mother), whereas any child penalty in same-sex couples is borne by one or both of the men in a same-sex couple.

Another factor to consider is whether to estimate weighted or unweighted regressions. Given the large number of likely non-random allocated values for labor-market outcomes, our preferred model is unweighted. The results for weighted regressions are in Appendix Fig. 4, using the "svy" command in Stata for person-level weights. The results are quite similar between the weighted and unweighted regressions. The sample weights are designed to make the sample representative of the U.S. population, not just the sample of people in couples. Although we cannot control for changes in cohabitation over time, the similarity of the weighted and unweighted results is consistent with there being no systematic changes in who chooses to cohabitate over time. Similarly, Hansen et al. (2020b) find no relationship over time between tolerance and the likelihood of being in a same-sex couple.

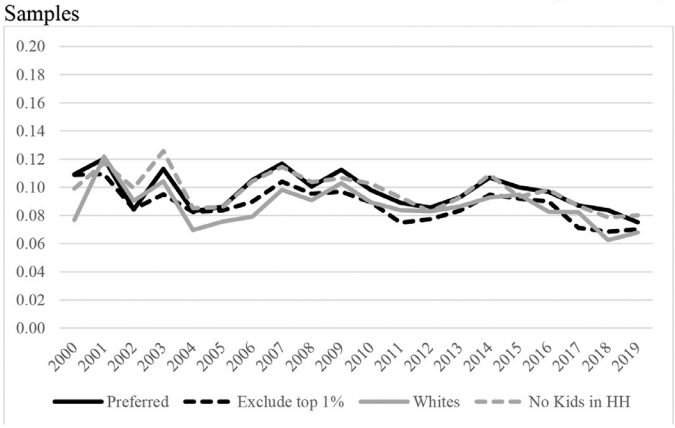
As mentioned previously, the results combine married and unmarried same-sex couples because the ACS does not identify married same-sex couples prior to 2012. In fact, prior to 2012, married same-sex couples were recoded as unmarried same-sex couples, and their marital status was allocated. Thus, they are excluded from our analysis because we cannot distinguish married same-sex couples from unmarried same-sex couples or even mis-classified different-sex couples among the individuals with allocated marital status. We have two reasons to believe that the trends found here, at least for 2008 to 2019 (after the survey redesign to prevent respondents from accidently marking both male and female genders), are unlikely to be affected by the exclusion of people with allocated marital status. First, Appendix Fig. 5 illustrates that the results from 2008 to 2019 are nearly identical when we include individuals with allocated marital status.[19] Second, the results for 2012 to 2014 are nearly identical when we exclude married same-sex couples. After 2014, the results differ as the number of married, same-sex couples increases dramatically. These results are available from the authors upon request.

[17] For brevity, the Figure does not include an additional sample whose results are nearly identical to the full sample where we exclude individuals in the fishing, forestry, and farming industries because almost no men or women in same-sex couples are employed in these industries.

[18] One possibility for this result, but one that we are unable to test, is that selection occurs if only economically successful lesbians have the confidence to be in a same-sex couple and 'come out' by choosing 'spouse' or 'same-sex partner' for the relationship to head of household.

[19] Prior to 2008, the results for same-sex couples with allocated marital status likely include a sizeable share of different-sex couples mis-coded as same-sex couples. Thus, the coefficients from 2000 to 2007 are sensitive to the inclusion or exclusion of couples with allocated marital status.

a: 2000-2019 ACS Coefficients for Women in Same-Sex Couples on Wages, Different Samples



b: 2000-2019 ACS Coefficients for Men in Same-Sex Couples on Wages, Different Samples

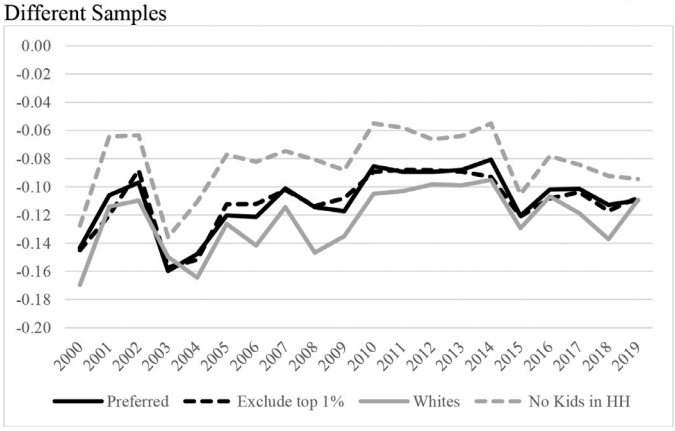


Fig. 9. Notes: Each point is the coefficient for individuals in same-sex couples (compared to married individuals in different-sex couples) from a separate regression, based on Eq. (1). The dependent variable is log annual wages. All coefficients are statistically different from zero at the one-percent level for a two-sided test.

In the results so far, we assume the relationship between sexual orientation and labor-market outcomes is the same for all individuals in same-sex couples, rather than varying with demographic characteristics. Martell (2020) shows that the wage premium for women in same-sex couples can be explained by differences in the return to experience. When he includes interaction terms between sexual orientation and potential experience, the coefficient for women in same-sex couples is negative. Given the positive interaction term between women in same-sex couples and potential experience, the wages of women in same-sex couples are higher than those of married women in different-sex couples for individuals with approximately 10 years of experience or more. We can replicate this pattern in our data throughout the time period, as expected, given that Martell (2020) also uses ACS data. A fruitful area for future research is to explore the role of other characteristics such as race / ethnicity and education in explaining differences in labor-market outcomes by sexual orientation.

One way to look at the role of changes in characteristics such as potential experience is to perform Oaxaca-Blinder decomposition.[20] In this decomposition, the raw gap in log wages between individuals in a same-sex couple and individuals in a different-sex couple is decomposed into the portion that can be explained by demographics and the remaining portion, labeled the 'unexplained' portion. Fig. 10 illustrates the decomposition. Log annual wages is the outcome, and the comparison group is the set of all different-sex couples (married or unmarried). By definition, the sum of the explained and unexplained portions is the size of the gap in log earnings.

For women, both the explained and unexplained portion of the wage gap declines substantially between 2000 and 2019. By 2017, the higher log wages for women in same-sex couples cannot be explained by their characteristics such as education. If anything, women in same-sex couples should have lower wages than those in different-sex couples if one only considers the portion of the wage gap explained by their characteristics.[21] The unexplained portion of the wage gap also declines substantially over time. By 2019, the unexplained portion has declined by half of its 2000 to 2003 value. As the unexplained portion contains the effects of discrimination, along with other factors, the decline in the unexplained portion is consistent with a decrease in discrimination.

For men, few if any patterns emerge in the explained and unexplained portions of the wage gap. The size of both gaps fluctuates over the time period, although the size of the explained portion appears to be smaller in 2017 to 2019 compared with previous years. The explained portion is negative, suggesting that men in same-sex couples have more favorable characteristics such as education that should lead to higher log wages, in contrast to the wage penalty observed in the coefficients for men in same-sex couples. In contrast to women, men in same-sex couples show no marked improvement in discrimination or other unexplained factors of the wage gap with married men in different-sex couples.

Although the labor-market outcomes of women in same-sex couples are generally better than those for married women in different-sex cou-

[20] The Oaxaca-Blinder decomposition is an extension of earlier work on differences across groups; Kitagawa (1955) summarizes early work on decompositions and provides an example of such a technique.

[21] One way to look at the effect of specific characteristics is to run a regression that contains only that characteristic, along with couple type, as controls. When looking at individual characteristics this way (results available upon request), we see that the regression-adjusted gap when we adjust only for education is similar to the regression-adjusted gap in Fig. 4 when we adjust for all the characteristics. In contrast, the regression-adjusted gap when controlling only for a different factor, such as race / ethnicity, is similar to the raw gap.

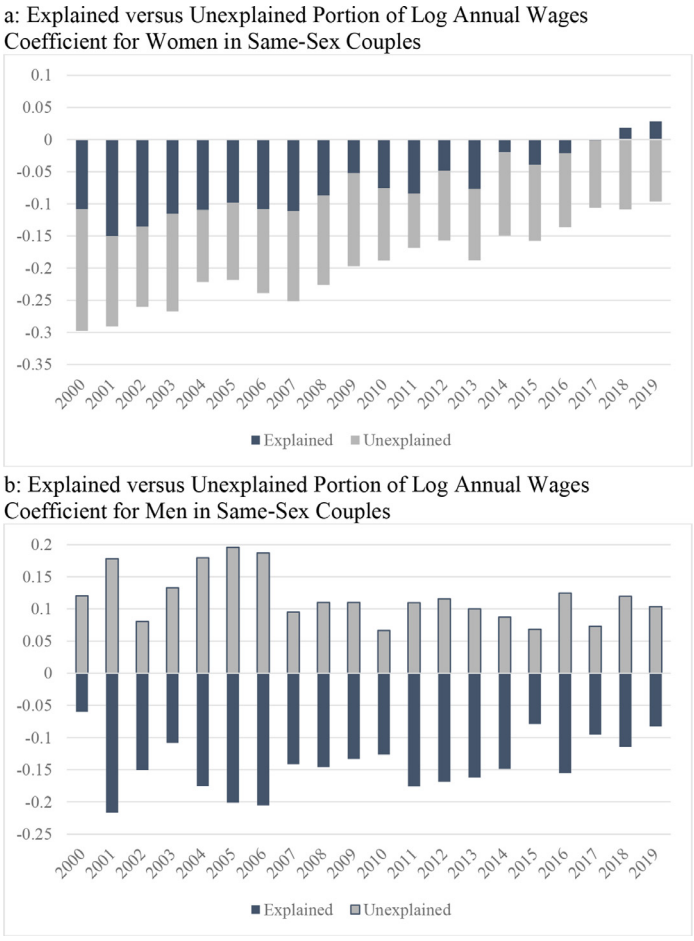


Fig. 10. Notes: Each point is the coefficient for individuals in same-sex couples, compared to individuals in different-sex couples (married or unmarried), from a separate regression, based on Eq. (1). The dependent variable is log annual wages. The bars represent the explained and unexplained portion of that coefficient from Oaxaca-Blinder decomposition.

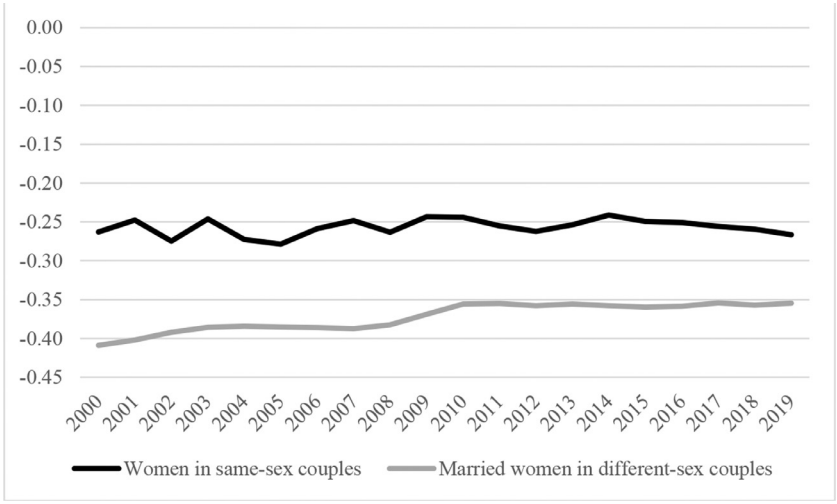


Fig. 11. Notes: Each point is the coefficient for either women in same-sex couples (black line) or married women in different-sex couples (gray line) based on Eq. (1). Men and women are pooled in the same regression, but separate regressions are estimated for each year. Although not shown in the Figure, the regression also includes couple type variables for men in same-sex couples, unmarried men in different sex couples, and unmarried women in different-sex couples. The dependent variable is log annual wages. All coefficients are statistically different from zero, and the coefficients for the two female couple types are statistically different from each other at the one-percent level for a two-sided test.

ples, these women may still suffer a gender wage gap compared to men. Thus, we estimate Eq. (1) on the pooled sample of men and women with the corresponding couple type variables. Fig. 11 contains the coefficients for women in same-sex couples and married women in different-sex couples. The omitted category is married men in different-sex couples, and the outcome is log annual wages.

Fig. 11 shows that the wages of women in same-sex couples are not converging toward those of men in different-sex couples. Instead, the wage penalty fluctuates between -0.25 and -0.28 log points, or 22–24%. In contrast, the gap between men and married women in different-sex couples has declined over the time period, from -0.41 log points (or 34%) in 2000 to -0.35 log points (or 30%) in 2019. At this rate of 0.054 log points per 20 years, convergence between married men and married women in different-sex couples would take 131 years.

Finally, changes in state and local policies and attitudes could also contribute to changes over time in labor-market differentials by sexual orientation. Several studies focus on state-level anti-discrimination laws (Klawitter and Flatt, 1998; Gates, 2009; Klawitter, 2011; Martell, 2013, 2014; Burn, 2018), and Delhommer (2020) also looks at local laws. Sansone (2019) and Hansen et al. (2020a) study same-sex marriage laws,

and Burn (2020) and Hansen et al. (2020b) explore state-level measures of tolerance. Our preferred model accounts for such state-level changes through the use of state fixed effects and separate regressions by year, but this method cannot directly measure the impact of these changes on labor-market outcomes.

6. Conclusion

Using 2000 to 2019 ACS data on cohabiting individuals, we find that the regression-adjusted gap in wages by sexual orientation narrows between 2001 and 2008. After that, the gap remains relatively flat for men in same-sex couples at around 11% for annual wages, earnings, and income. The premium for women in same-sex couples declines slightly in later years, with a gap of around eight percent in 2018. The rate of decline is sufficiently slow, however, so that convergence in wages would not be reached for at least 25 years if the current trend in convergence continues. In contrast, the raw gap in wages between women in same-sex couples and married women in different-sex couples has declined dramatically between 2000 and 2019.

Although our analysis provides insight on earnings patterns by sexual orientation, the study is descriptive rather than causal. ACS data only identify individuals in same-sex couples, so the analysis is limited to members of same-sex couples. Although many researchers document the recent increase in the share of different-sex couples who prefer unmarried cohabitation over married cohabitation (Manning, 2020; Manning et al., 2019), there is little evidence of dramatic changes in couples' decision to cohabitate. For example, Carpenter (2020) does not find a significant effect of the legalization of same-sex marriage on the cohabitation decision of same-sex couples. More generally, changes over time in cohabitation, particularly among same-sex couples, is an important topic deserving of further study.

Using ACS data, we do not find evidence of convergence of wages for men, which is in contrast to recent work suggesting that the convergence is continuing (in Canada: Dilmaghani, 2017; Mueller, 2014), has already converged (in the UK: Aksoy et al., 2018), or that gay men actually earn more than heterosexual men in the U.S. National Health Interview Survey (Carpenter and Eppink, 2017). Aksoy et al. (2018) document a difference by cohabitation in the UK. For cohabiting couples, their results match ours; for individuals who are not cohabiting, earnings are not statistically different by sexual orientation in their sample.

We seek a better understanding of the discrepancies in earnings by sexual orientation. We are limited in our ability to analyze earnings across living arrangements when using ACS data. The largest data set with detailed information on sexual orientation and labor-market outcomes is the National Health Insurance Survey (NHIS). The NHIS includes roughly 35,000 individuals per year, yet the sample sizes for gay men and lesbians are quite small. Many of the subgroup analyses in Carpenter and Eppink (2017) produced statistically insignificant effects due to small sample sizes. Thus, we echo their demand for the inclusion of sexual orientation information in large data sets such as the ACS.

Declaration of Competing Interest

None.

Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.labeco.2021.102086.

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