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# Fraternal Birth Order, Only-Child Status, and Sibling Sex Ratio Related to Sexual Orientation in the Add Health Data: A Re-Analysis and Extended Findings

**Malvina N. Skorska, Ph.D.**[1,2], **Anthony F. Bogaert, Ph.D.**[2,3]

[1]Department of Psychology, University of Toronto Mississauga, 3359 Mississauga Road North, Mississauga ON L5L 1C6

[2]Department of Psychology, Brock University, St. Catharines ON

[3]Department of Health Sciences, Brock University, St. Catharines ON

## Abstract

The fraternal birth order (FBO) effect related to men's sexual orientation refers to the finding that the number of older brothers that a man has increases his chance of being androphilic. The FBO effect has generally been well replicated in diverse samples; one instance of non-replication was by Francis (2008) using Waves I and III of the Add Health data. We attempted to replicate the FBO effect in the Add Health data taking into account family size and other limitations of Francis's (2008) analyses. Also, we examined other sibling characteristics related to the FBO effect: sibling sex ratio and only-child status. We used two subsamples from Waves I ($n$ = 20,745), and IV ($n$ = 15,701) of the Add Health data, consisting of adolescents who were followed longitudinally from 1994-1995 until 2008. Wave I data was used to compute numbers of younger and older brothers and sisters from household roster information. Wave IV information about sexual orientation identity was used. Analyses were conducted within men and within women. We found modest support for the FBO effect in men, but not in women, using the older brother odds ratio, logistic regression analyses, and sibling sex ratio, which provided the strongest support for FBO. We found that gynephilic/biphilic women, but not androphilic/biphilic men, were more likely to be only-children compared to androphilic women and gynephilic men, respectively. We discuss limitations of the Add Health data and purported mechanisms for the FBO effect in men and the only-child effect in women.

### Keywords

sexual orientation; fraternal birth order; sibling sex ratio; only-child status; Add Health

**Corresponding Author:** Malvina N. Skorska, Department of Psychology, University of Toronto Mississauga, 3359 Mississauga Road North, Mississauga ON L5L 1C6, Phone: 905-569-5789, malvina.skorska@utoronto.ca.

COMPLIANCE WITH ETHICAL STANDARDS

The authors declare no conflict of interest.

All procedures performed in studies involving human participants were in accordance with the ethical standards of the institutional and/or national research committee and with the 1964 Helsinki declaration and its later amendments or comparable ethical standards.

Informed consent was obtained from all individual participants included in the study by the Add Health researchers. Ethics approval for secondary data analysis was received from the University Research Ethics Board.

## INTRODUCTION

The fraternal birth order (FBO) effect related to men's sexual orientation refers to the finding that the number of older brothers that a man has increases his chance of being androphilic (i.e., attracted to men). Blanchard (2018a) conducted a meta-analysis of 26 studies published between 1992 and 2017 testing the FBO effect in 30 androphilic male groups and 30 gynephilic (i.e., attracted to women) male groups. The older brother odds ratio (OBOR) was established to provide a standard metric to compare the number of older brothers between androphilic and gynephilic men across the samples independent of family size. An OBOR of 1 suggests no difference, and a value greater than 1 suggests androphilic men have greater numbers of older brothers than gynephilic men. An OBOR of 1.47 was found and it was significantly different from 1. When the androphilic male group was further separated into a feminine/transgender androphilic sample and a masculine androphilic sample, the OBOR was 1.85 for the former and 1.27 for the latter, with both OBORs significantly different from 1 and significantly different from each other. Family size was further ruled out by showing that the OBOR did not differ between groups who had smaller versus larger family sizes. Thus, Blanchard (2018a) demonstrated that androphilic men had a greater number of older brothers than gynephilic men.

Blanchard (2018a) originally excluded studies using probability samples because of imprecise measures of siblings, but in a follow-up meta-analysis, Blanchard (2018c) demonstrated that the OBOR was 1.21 and was significantly greater than 1 in studies using these probability samples. Blanchard (2018b) further showed that older sisters have no effect or a weaker effect compared to older brothers using the same samples as Blanchard (2018a). Across the studies included in the meta-analyses, it has also been demonstrated that year of birth, number of older sisters, number of younger brothers, number of younger sisters, age of parents at time of birth, age of participants, socioeconomic status, and sibship size were not confounds to the FBO effect (see Blanchard, 1997 and Bogaert & Skorska, 2011 for reviews). Further studies have indicated that the FBO effect can explain the sexual orientation of about 15-29% of androphilic men (Blanchard & Bogaert, 2004; Cantor, Blanchard, Paterson, & Bogaert, 2002).

Since the publication of Blanchard's meta-analyses, five studies have been published further supporting the FBO effect. Li and Wong (2018) showed that androphilic men had a greater number of older brothers than gynephilic men in Hong Kong, and the effect was not observed for gynephilic women (vs. androphilic women). Apostolou (2019) also found that androphilic men had a greater number of older brothers than gynephilic men in a Greek sample (but not gynephilic women vs. androphilic women). Coome, Skorska, van der Miesen, Peragine, and VanderLaan (2018) showed that boys who were gender variant had a greater number of older brothers than boys who were not gender variant. Birth order was not associated with gender variance in girls. Similarly, androphilic male adolescents had a greater number of older brothers than biphilic and gynephilic male adolescents, whereas number of older brothers was not associated with sexual orientation in female adolescents (Xu, Norton, & Rahman, 2019). Swift-Gallant, Coome, Monks, and VanderLaan (2018) and Wampold (2018) showed that androphilic men who were bottoms in their anal sex role had a

greater number of older brothers compared to androphilic men who were not bottoms (both studies) and gynephilic men (Swift-Gallant et al., 2018 only). Overall, regarding men's sexual orientation, there is a strong body of evidence suggesting that androphilic men have a greater number of older brothers than gynephilic men.

The most developed explanation for the FBO effect is biological in nature. Specifically, a maternal immune hypothesis (MIH) has been forwarded as the most likely explanation, where a mother's immune response, including a build-up of antibodies, to a male-specific protein increases with each successive male pregnancy (Blanchard & Bogaert, 1996). The increasing amounts of antibodies would cross the blood-brain barrier to affect sites in the brain responsible for sexual orientation in a later-born developing fetus. The biological basis of the FBO effect has received support through the replication of this effect in various countries with diverse cultures (e.g., Brazil, Iran, Samoa, United States, etc.; see Blanchard, 2018a). Bogaert (2006) also found direct support for the biological basis of the FBO effect by demonstrating that only biological older brothers were significantly related to androphilia in men, but not biological younger brothers, biological older sisters, biological younger sisters, and nonbiological siblings (older or younger brothers or sisters). The MIH itself received direct support in 2018 with a study (Bogaert et al., 2018) examining antibody concentrations to NLGN4Y, a cell-surface protein expressed in the male brain (Jamain et al., 2003; Skaletsky et al., 2003). Controlling for total numbers of pregnancies, mothers of androphilic men with older brothers had the highest anti-NLGN4Y concentrations compared to mothers of androphilic men without older brothers, mothers of gynephilic men, and women with no sons (Bogaert et al., 2018). Thus, Bogaert et al. (2018) provided the first direct empirical evidence for the MIH underlying the FBO effect and an immunological influence on the development of sexual orientation in men.

In women with gynephilia, early studies found mixed results (early, late or no relationship to birth order; see Blanchard, 1997; Bogaert, 1997), and in newer studies, gynephilic women have generally not been found to have a greater number of older brothers or older sisters (Apostolou, 2019; Blanchard et al., 1998; Bogaert, 2003; Coome et al., 2018; Li & Wong, 2018; Xu et al., 2019). Thus, birth order does not seem to be reliably associated with women's sexual orientation. However, a sibling characteristic that has been associated with sexual orientation in women is only-child status. VanderLaan, Blanchard, Wood, and Zucker (2014) found that gynephilic adolescent girls who experience gender dysphoria were more likely to be only-children compared to androphilic and gynephilic adolescent boys (an androphilic adolescent girl sample was not included). Similarly, Schagen, Delemarre-van de Waal, Blanchard, and Cohen-Kettenis (2012) found that adolescent girls who experience gender dysphoria were more likely to be only-children compared to control girls, whereas there was no difference between adolescent boys who experience gender dysphoria and control boys.

Blanchard (2012b) hypothesized that an immunological effect independent of the FBO effect could explain a greater likelihood of only-children with a non-heterosexual sexual orientation. Specifically, studies have shown that a maternal immune response to a fetus is associated with a reduced weight in the newborn and with increased miscarriages in later pregnancies (Kahn & Baltimore, 2010; Nielson, 2011). Support for this independent

immune response was found in Blanchard (2012a) by showing that seven gynephilic women and six androphilic men who were only-children had a lower birth weight than androphilic women and gynephilic men who were only-children and all participants who were the oldest child. However, VanderLaan, Blanchard, Wood, Garzon, and Zucker (2015) only replicated the birth weight finding in male individuals referred for gender dysphoria and not in female individuals referred for gender dysphoria. Skorska, Blanchard, VanderLaan, Zucker, and Bogaert (2017) showed a lower birth weight in a sample of first-born androphilic men who were only-children, and mothers of androphilic men who were only-children had more miscarriages than mothers of men with other sibship compositions (women were not investigated). Thus, there is some support for the finding that gynephilic women might be more likely to be only-children, but the results sometimes extend to androphilic men and further research is necessary using samples of both men and women.

Sibling sex ratio (SSR) is another family biodemographic marker that has been examined with respect to sexual orientation. SSR is the ratio of brothers to sisters of an individual (i.e., male:female siblings) and is estimated to be 106:100 in the general population (James, 1987; Orzack et al., 2015). Blanchard (1997) found that the SSR of male-to-female androphilic transsexuals was 134:100, and it was significantly greater than 106:100. It was driven by an excess of both older and younger male siblings. For more masculine androphilic men, SSR was 108:100, and it did not differ from 106:100; however, the SSR was about 110:100 for older siblings and about 99:100 for younger siblings. VanderLaan et al. (2014) found that androphilic boys who experience gender dysphoria had a SSR of 128:100, which was driven by an older SSR (i.e., the older SSR was 136:100 and the younger sibling sex ratio was 119:100). The SSR for more masculine androphilic men has been interpreted as a reflection of the FBO effect. For the feminine or transgender androphilic women, it was proposed that a dosage effect related to a maternal immune response was occurring (Blanchard 1997; see also Blanchard 2018a). That is, older brothers might result in a smaller/weaker maternal immune response, only affecting the sexual orientation of men. However, having an excess of siblings might result in an even larger maternal immune response (and likely one that differs from the immune response purported to explain the only-child effect), affecting both sexual orientation of offspring and emotional and behavioral characteristics to be more feminine.

Thus, in addition to SSR being a separate biodemographic variable worthy of examining in relation to sexual orientation, it potentially provides insight about mechanisms underlying the association between FBO and sexual orientation. Moreover, SSR might also provide insights regarding the sample utilized in the study given knowledge of a population SSR. For example, if a sample (e.g., comprising both gynephilic and androphilic participants) has a higher SSR than the estimate of the SSR in the general population, then that sample might be unusual in that it over-represents the number of boys born to the mothers in that sample.

In the current study, we examined the FBO effect, only-child status, and SSR using Waves I and IV of the National Longitudinal Study of Adolescent to Adult Health (Add Health; Harris, 2009). The goal of Add Health was to collect data about the health of a cohort of American adolescents in grades seven to twelve (Harris et al., 2009). Add Health consists of a nationally representative sample of adolescents and young adults followed from 1994 until

2008 with four waves of data collection. At Wave I in 1994, participants were 11 to 18 years of age and at Wave IV in 2008, the majority of participants were between 24 and 34 years of age. The Add Health data was used in one previous study of the FBO effect and sexual orientation by Francis (2008).

Francis (2008) used Waves I, II, and III of the Add Health data. Specifically, a behavioral measure of sexual orientation was derived from Waves I, II, and III. Romantic attraction was derived from Wave III and sexual orientation identity was derived from Wave III. Sibling variables were constructed using information about who was living with the adolescent at the time of Wave I and from a question that asked how many children a participant's biological parents had together, again at Wave I. Within male adolescents, Francis (2008) found that multiple older sisters was related to a gynephilic sexual orientation, and noted that multiple older brothers was associated with an androphilic sexual orientation and would likely be significant at $p < .10$. Within female adolescents, one older brother, one younger sister, and one older sister were generally significantly associated with an androphilic sexual orientation. Blanchard (2014) proposed that a potential reason for the non-replication by Francis (2008) was not taking into account family size. Indeed, Blanchard (2014) demonstrated that the FBO effect is more likely to be found when controlling for total family size.

In addition to Blanchard's (2014) concern, we note the following limitations with Francis' (2008) analyses. First, it is unclear whether Francis conducted separate analyses for all sibling variables or entered them all into one regression analysis, which could pose a problem given sibling variables are generally related. Furthermore, most analyses of the FBO effect prioritize older brothers given the theoretical significance of this variable. Second, Francis included participants with same-age siblings (e.g., twins) in analyses, which Blanchard (2018c) indicated were excluded in his initial meta-analysis on theoretical grounds. That is, it would be difficult to determine which sibling is considered "older" and it is difficult to make predictions about the uterine environment of each same-age sibling. Third, Francis (2008) categorized FBO (i.e., one older brother vs. multiple older brothers), which Blanchard (1997) noted is a limitation of early FBO research because it decreases variability associated with numbers of siblings. Fourth, Francis (2008) did not provide details on how the question about number of children a participant's biological parents had together at Wave I was used with the household roster information to provide estimates of each category of siblings (see Methods for more details regarding complications associated with this question). Fifth, Wave IV sexual orientation data could not be used (it was not available at that time). Savin-Williams and Joyner (2014) noted potential problems with using measures of romantic attraction completed by participants at Wave I and suggest using Wave III and/or IV sexual orientation identity data (cf. Katz-Wise, Calzo, Li, & Pollitt, 2015; Li, Katz-Wise, & Calzo, 2014). Although recently it was shown there is no impact on health disparity findings using earlier waves and romantic attraction measures in Add Health (Fish & Russell, 2018), using Wave IV sexual orientation identity might result in different findings given changes in sexual orientation shown in the Add Health data especially in women (Savin-Williams, Joyner, & Rieger, 2012). Furthermore, we are interested in adult sexual orientation; therefore, extending the work of Francis (2008) to examine Wave IV is pertinent.

Thus, the present study is an attempted replication and extension of Francis (2008), where Wave IV sexual orientation identity is used, FBO is not a categorical variable, participants who have same-age brothers are excluded, family size is accounted for, older brothers are prioritized in regression analyses, and Blanchard's OBOR is computed to place the findings within the context of the results of the meta-analysis. Note that the results of the OBOR for one subsample reported in our study here (see also Skorska & Bogaert, 2017b) were reported in Blanchard (2018c). Here we provide details on the computation of this value and also expand on the finding reported in Skorska and Bogaert (2017b) and Blanchard (2018c). We also computed SSR to provide additional information about the FBO effect and provide population-level comparisons of SSR estimates. Last, we examined whether only-child status is associated with sexual orientation. All analyses were conducted separately within men and within women.

We predicted that replication of the FBO effect in men will occur in the Add Health data. In women, we did not predict a relationship with older or younger brothers or sisters and sexual orientation. We predicted that androphilic men will have a greater SSR than gynephilic men and it will be driven by an older SSR. In women, we predicted a lower SSR in gynephilic women compared to androphilic women based on the results of VanderLaan et al. (2014). We predicted that gynephilic women and androphilic men will be more likely to be only-children compared to androphilic women and gynephilic men based on the findings of Blanchard (2012a), Schagen et al. (2012), Skorska et al. (2017), VanderLaan et al. (2014), and VanderLaan et al. (2015).

## METHOD

### Participants

At Wave I, 20,745 adolescents were interviewed at home (see Procedure section for further details). At Wave IV 15,701 of the original Wave I respondents were re-interviewed at home. We used two sub-samples in analyses. One subsample, hereafter named Subsample I, of 19,820 adolescents is the largest sample including Wave I participants that does not include those who had a same-age brother (presumably a co-twin, which obscures birth order position among the male participants) and those who had inconsistent sex across Waves I and IV. This subsample was used to maximize the sample of androphilic males and gynephilic females, and thus maximize statistical power. This subsample was also included in Blanchard (2018c) in the meta-analysis of the OBOR in probability samples. We also used a subsample, hereafter named Subsample II, of 14,172 adolescents who had a Wave IV cross-sectional sample weight (see Procedure section for further details; see also Chen & Chantala, 2014; Savin-Williams & Joyner, 2014), but did not include participants with same-age brothers or inconsistent sex. Not all analyses contain data for all participants within each subsample and analysis-relevant *n*'s are reported throughout. Codebooks for the Add Health data are available online ("Questionnaire Codebooks for Waves I, II, III and IV," n.d.).

For Subsample I, the mean age of men at Wave IV was 28.65 ($SD = 1.79$, $n = 6883$) and the mean age of women at Wave IV was 28.43 ($SD = 1.78$, $n = 8168$). For Subsample II, the mean age of men at Wave IV was 28.67 ($SD = 1.78$, $n = 6528$) and the mean age of women at Wave IV was 28.45 ($SD = 1.78$, $n = 7711$). For Subsample I, 57.6% ($n = 5572$) of men

were White and 57.4% ($n$ = 5898) of women were White. Those of other ethnicities comprised 42.2% ($n$ = 4081) of the sample in men and 42.4% ($n$ = 4350) of the sample in women. For Subsample II, 60.7% ($n$ = 3962) of men were White and 59.3% ($n$ = 4574) of women were White. Those of other ethnicities comprised 39.2% ($n$ = 2557) of the sample in men and 40.5% ($n$ = 3123) of the sample in women. For Subsample I, 43.0% ($n$ = 4155) of men had a university education and 55.8% ($n$ = 5728) of women had a university education. Those with other education comprised 28.2% ($n$ = 2726) of the sample in men and 23.7% ($n$ = 2438) of the sample in women. For Subsample II, 60.9% ($n$ = 3975) of men had a university education and 70.5% ($n$ = 5433) of women had a university education. Those with other education comprised 39.1% ($n$ = 2551) of the sample in men and 29.5% ($n$ = 2276) of the sample in women.

## Measures

Only the measures of interest to the current study are described below. Any responses in which the participant selected “refused,” “don’t know,” “not applicable” or “legitimate skip” were coded as missing data.

**Sex (Waves I and IV).—**Biological sex was self-reported by the Add Health interviewers as either male or female. Sex that was consistent across Waves I and IV was used. For some participants, only Wave I sex was available and this information was used to determine the sex of the participant.

**Age (Wave IV).—**Age at Wave IV was provided as part of the Add Health data. It was calculated by Add Health using the date of the interview and date of birth reported by participants.

**Race/Ethnicity (Wave I).—**The race/ethnicity variable was dichotomized as White or other ethnicities (as in Skorska & Bogaert, 2017a). Participants could choose more than one category to represent their race/ethnicity. Those who selected only the White category were coded as 0, and those who selected any combination of other ethnicities were coded as 1.

**Education (Wave IV).—**The education variable was dichotomized as completion of at least some university/college education (coded 1) and no completion of any university/college education (coded 0) (as in Skorska & Bogaert, 2017a). Specifically, those who indicated they completed “eighth grade or less,” completed “some high school,” were a “high school graduate,” completed “some vocational/technical training (after high school),” or “completed vocational/technical training (after high school)” were coded 0 and all other participants who answered the education question were coded 1.

**Sexual orientation (Wave IV).—**At Wave IV, participants were asked about their sexual orientation identity via the following statement, “Please choose the description that best fits how you think about yourself.” The response options were: “100% heterosexual (straight),” “mostly heterosexual (straight), but somewhat attracted to people of your own sex,” “bisexual that is, attracted to men and women equally,” “mostly homosexual (gay), but somewhat attracted to people of the opposite sex,” “100% homosexual (gay),” and “not

sexually attracted to either males or females." Responses to the "not sexually attracted to either males or females" option were coded as missing data for the current study.

The main groups of interest were the predominantly gynephilic men and androphilic women (coded 0; for Subsample I, $n_{men}$ = 6,598, $n_{women}$ = 7,759; for Subsample II, $n_{men}$ = 6,262, $n_{women}$ = 7,325) compared to the predominantly androphilic men and gynephilic women (coded 1; for Subsample I, $n_{men}$ = 174, $n_{women}$ = 144; for Subsample II, $n_{men}$ = 167, $n_{women}$ = 137). One other grouping was also of interest: predominantly gynephilic men and androphilic women (coded 0; for Subsample I, $n_{men}$ = 6,598, $n_{women}$ = 7,759; for Subsample II, $n_{men}$ = 6,262, $n_{women}$ = 7,325) versus other sexual orientation (coded 1; for Subsample I, $n_{men}$ = 225, $n_{women}$ = 337; for Subsample II, $n_{men}$ = 212, $n_{women}$ = 319). Predominantly gynephilic men and androphilic women were those who selected either "100% heterosexual (straight)" or "mostly heterosexual (straight)"; predominantly androphilic men and gynephilic women were those who selected either "100% homosexual (gay)" or "mostly homosexual (gay)"; and others were those who selected "bisexual," "mostly homosexual (gay)", or "100% homosexual (gay)."

**Numbers of older brothers, older sisters, younger brothers, and younger sisters (Wave I).—**Numbers of full older brothers, full older sisters, full younger brothers, and full younger sisters were taken from household roster information participants completed at Wave I, similar to Francis (2008). Specifically, participants reported the individuals that were currently living in the same household as them, the household member's sex, and the participant's relationship to the household member. If participants answered that a household member was a brother or sister, the participant then chose whether that household member was a full, half, step, adoptive, foster, or other sibling. Participants also provided the age of the individual (and other information not necessary for this study). We subtracted age of the household member from the participant's age to determine whether the household member was older, younger, or the same age. From this information we were able to count the number of full, half, step, adoptive, foster, or other brothers and sisters that were older, younger, or the same age for each participant.

As mentioned in the Participants section, we used two sub-samples and both subsamples filtered out participants who had a same-age brother. Given use of a household roster, we reasoned that any half-siblings would likely be maternal in origin since individuals tend to live with their mothers following a separation between a mother and father, and especially in the 1990s (Kelly, 2007; Meyer, Cancian, & Cook, 2017). Also, given FBO is a proxy for the MIH, we reasoned that we could include participants who had same-age sisters, who should theoretically not have an impact on generation of antibodies to male-specific proteins. Of course, this logic would not apply to same-age brothers, which is why participants with same-age brothers were excluded. Participants with step, foster, adoptive, or other siblings were not excluded, but totals of these types of siblings were not included in calculations.

At Wave I there were additional questions about biological children and birth order related to biological children. The specific questions asked, "How many children have your biological parents had together?" and "Which child are you—the first, the second, or what?" Francis (2008) used these questions, but did not provide details on how these questions were

used with the household roster information to calculate estimates of each category of siblings. Given that these questions do not provide information about the sex of biological children, we decided to not use these questions.[1]

Thus, from the household roster at Wave I, we used numbers of full and half brothers and sisters that were older or younger, as well as numbers of same-age sisters. From these variables, we could compute the following variables:

**1) Older brother odds ratio (OBOR).:** The total number of each relevant category of siblings was counted for each subsample. As per Blanchard (2018a), the OBOR was calculated as:

OBOR = (androphilic men's number of older brothers ÷ androphilic men's number of other siblings) ÷ (gynephilic men's number of older brothers ÷ gynephilic men's number of other siblings).

**2) Modified proportion of older brothers, modified proportion of older sisters, modified proportion of younger brothers, modified proportion of younger sisters, and modified proportion of same-age sisters.:** For each participant, variables representing full and half older brothers, older sisters, younger brothers, younger sisters, and same-age sisters were calculated based on numbers of each of these sibling categories and number of total siblings. These variables take into account family size and are able to utilize participants who are only-children to maximize statistical power. These variables were calculated as per Blanchard (2014):

Modified proportion of older brothers = (older brothers + 0.25)/(total siblings + 1),

Modified proportion of older sisters = (older sisters + 0.25)/(total siblings + 1),

Modified proportion of younger brothers = (younger brothers + 0.25)/(total siblings + 1),

Modified proportion of younger sisters = (younger sisters + 0.25)/(total siblings + 1), and Modified proportion of same-age sisters = (same-age sisters + 0.25)/(total siblings + 1).

**3) Sibling sex ratio (SSR).:** The total number of each relevant category of siblings was counted for each subsample. As per VanderLaan et al. (2014) the relevant variables were calculated as:

SSR = brothers/sisters*100,

Older SSR = older brothers/older sisters*100, and

Younger SSR = younger brothers/younger sisters*100.

[1]Also, at Wave IV a question was asked about total number of siblings: "How many brothers and sisters do you have, both living and deceased? Include biologically related, adoptive, and step-brothers or sisters." We decided not to use this question because it provides no information about biological siblings, sex of the siblings, or age of the siblings. A household roster, similar to the Wave I household roster, was also completed at Wave IV. We decided not to use the Wave IV household roster given many individuals at that age (between 24 to 34 years of age) did not live with their siblings

Same-age sisters were excluded in the older and younger SSR calculations because it was unclear where they should be included.

**4) Only-child status.:** To determine whether a participant was an only child, a total siblings variable was calculated for each participant. If a participant had 0 total siblings, they were considered an only-child (coded 1). If a participant had 1 or more siblings, they were not an only-child (coded 0).

## Procedure

Full details about the school-based sampling design can be found in Harris (2013) and Skorska and Bogaert (2017a). In 1994 and 1995 (Wave I), 90,118 students completed a short in-school questionnaire and a subsample of students was chosen to complete an in-depth 1.5-hour in-home interview. The Wave I in-home sample contains responses from 20,745 adolescents including special supplemental subsamples (e.g., students with certain ethnicities were oversampled). The in-home sample that does not include special supplemental subsamples is a representative sample of US adolescents in grades seven to 12. Interviews at Waves II, III, and IV were based on the Wave I sample of adolescents who completed in-home interviews. At Wave IV, 15,701 of the original Wave I in-home participants were re-interviewed in a 90-minute in-home interview.

## Statistical Analyses

All analyses were conducted using SPSS (version 22). Generally, analyses using the Add Health data should be conducted taking the sampling strategy into account (e.g., using Complex Samples in SPSS). Where Complex Samples was used in the current study, the appropriate Wave IV cross-sectional weight, stratum variable, and cluster variable were utilized to provide correct estimates of totals, ratios, regression parameters, means, variances, standard errors, and confidence intervals (Chen & Chantala, 2014). A “with replacement” design type was specified and subpopulation variables were created and utilized for analyses in which only a subset of the sample was to be analyzed (Chen & Chantala, 2014). A common subpopulation variable that was used was sex, because all analyses were conducted within each sex. Some of the analyses do not use Complex Samples because some of the calculations involved frequency counts and frequency counts cannot be calculated using Complex Samples. Subsample II, however, is a subsample that only includes participants with a Wave IV cross-sectional weight in an attempt to somewhat adjust for not using Complex Samples.

Analyses with OBOR and SSR used frequency counts and thus did not use Complex Samples. For OBOR, an online calculator (https://select-statistics.co.uk/calculators/confidence-interval-calculator-odds-ratio/; “Odds-ratio—confidence interval calculator,” 2019) was used to calculate 90% and 95% confidence intervals (CI) and to test whether the OBOR was significantly greater than 1. If the 90% and 95% CI did not include 1, the OBOR was significant at $p = .10$ and $p = .05$, respectively. In the contingency table, the numbers of full and half older brothers and numbers of total other siblings (older sister + younger brothers + younger sisters + same-age sisters) for gynephilic men and androphilic men were inputted. The same procedure was used to compare androphilic women and gynephilic

women. A different online calculator (https://www.medcalc.org/calc/odds_ratio.php; "Odds ratio calculator," 2019) was used to provide the two-tailed *p*-values associated with the 95% CI.

For SSR, the binomial probabilities online calculator (http://vassarstats.net/index.html; Lowry, 2019b) was used to compare the SSR for each group to the population SSR (106:100 male:female births) using the *z*-approximation to the binomial test. Values have to be inputted for *n*, *k*, and *p* in the online calculator. For SSR, *n* = total number of siblings, *k* = total number of brothers, and *p* = 0.5146, the expected proportion in the population (i.e., *p* = odds/[1 + odds] = 1.06/[1+ 2.06] = 0.5146). For older SSR, *n* = total number of older siblings, *k* = total number of older brothers, and *p* = 0.5146. For younger SSR, *n* = total number of younger siblings, *k* = total number of younger brothers, and *p* = 0.5146. To test for differences between gynephilic and androphilic men, an online calculator (http://vassarstats.net/propdiff_ind.html; Lowry, 2019a) examined the difference between two independent proportions using the *z*-test for independent proportions.

For testing the modified proportion of older brothers and other sibling variables, logistic regression analyses were conducted both not using and using Complex Samples. Because of the nature of OBOR and SSR, analyses using these ratios could not be done using Complex Samples. The dependent variable was sexual orientation. In analyses that did not use Complex Samples, the following variables were directly entered on step 1: age (Wave IV), education, race/ethnicity, modified proportion of older brothers, and modified proportion of younger brothers. Given multi-collinearity concerns with the modified proportions variables, the following variables were entered in a forward stepwise entry (i.e., they would only enter the model if significant) on step 2: modified proportion of younger sisters, modified proportion of older sisters, and modified proportion of same-age sisters. Older brothers and younger brothers were entered together on Step 1 given the findings with SSR (see Results) and given these variables were correlated. In Complex Samples analyses, the dependent variable was sexual orientation and the independent variables were age, education, ethnicity, modified proportion of older brothers, and modified proportion of younger brothers, given multi-step analyses could not be run using Complex Samples.

For only-child analyses, logistic regression analyses were conducted both not using and using Complex Samples. The dependent variable was sexual orientation. In both sets of analyses, the following variables were directly entered on step 1: age (Wave IV), education, race/ethnicity, and only-child status. Age, education, and race/ethnicity were statistically controlled because of their relationship to some or all of the sibling variables, only-child status, and sexual orientation (results not shown). Also, for Subsample I, certain ethnicities and education levels were over-sampled (e.g., Black adolescents with highly educated parents; Harris, 2013) and thus controlling for these variables helps to take into account the sampling strategy. Results for the control variables of age, education, and ethnicity should be interpreted with caution given the direction of findings not using Complex Samples sometimes was opposite to what was found with Complex Samples. Perhaps this effect (including a change in direction) is not surprising given the weighting procedures associated with Complex Samples specifically targets (and thus alters) some of these demographic variables. Notably, the direction of the findings for older brothers and younger brothers,

however, were consistent across analyses. Thus, in keeping with our main emphasis on the sibling variables, our comments in the Results focus on these sibling variables (e.g., older brothers variables) and we have chosen not to comment on the age, education, and ethnicity results per se (but see Tables 6 and 7).

## RESULTS

### Older Brother Odds Ratio (OBOR)

Numbers of each of the sibling categories are found in Table 1 for men and Table 2 for women. OBOR results can be found in Table 3 for men and for women. In men, the OBOR ranged from 1.19 to 1.29 across the samples. Using one-tailed tests, one of these tests was significant at $\alpha = .05$ and one is approaching significance. However, the 95% CI, using a two-tailed test always included 1 and was not significant. In two of four cases the lower bound was > .90. In two of four remaining cases, the lower bound was > .88. For the 90% CI one lower bound was at 1 and all are > .90. Thus, the OBOR was significantly or marginally significantly different from 1 using a one-tailed test in two instances. Some modest support for androphilic/biphilic men being more likely to have older brothers than gynephilic men was found.

In women, the OBOR ranged from 0.98 to 1.13 across the samples and the lower bound does not overlap with the range of lower bounds for the OBOR in men. All 95% and 90% CIs include 1, and are nearer to 1 than the OBORs and CIs for men. The 95% CI are not significant with a two-tailed or one-tailed test. Gynephilic/biphilic women were not more likely to have older brothers than androphilic women.

### Sibling Sex Ratio (SSR)

Table 4 shows SSR by sexual orientation, and comparisons to the population value of 106:100 in men and in women. Table 5 shows comparisons between gynephilic and androphilic men, between gynephilic men and androphilic/biphilic men, between androphilic and gynephilic women, and between androphilic and gynephilic/biphilic women. In gynephilic men, the SSR was 108:100 and it was not significantly different from the population value (Table 4). In androphilic/biphilic men, the SSR ranged from 134:100-138:100 and was significantly or marginally significantly different from the population value using the standard two-tailed ($p < .05$) test. It would be significant in all cases with a one-tailed test.

To determine if older or younger brothers are driving the SSR results, the older SSR and younger SSR were computed. For both gynephilic and androphilic/biphilic men, the younger SSR was not significantly different from the population value. The older SSR for gynephilic men ranged from 115:100-116:100 and was significantly different from the population value. The older SSR for androphilic/biphilic men ranged from 174:100-197:100 and was significantly different from the population value. Last, the difference between gynephilic men and androphilic/biphilic men in SSR was examined (Table 5). For the SSR, in all cases gynephilic men had a marginally significantly lower SSR than androphilic/biphilic men and this finding would be significant with a one-tailed test. For the younger SSR, there were no

differences between gynephilic men and androphilic/biphilic men. For the older SSR, gynephilic men had marginally significantly or significantly lower older SSR than androphilic/biphilic men, and these findings would all be significant with a one-tailed test. Thus, androphilic/biphilic men tended to have a higher SSR, which was largely driven by the older SSR, compared to the population value and compared to gynephilic men.

In androphilic women, the SSR ranged from 94:100-96:100 and in both cases, it was significantly different from the population value, but in the opposite direction compared to the population value and compared to men: there were less male siblings than expected (Table 4). In gynephilic/biphilic women, the SSR ranged from 95:100-105:100, but these were not significantly or marginally significantly different from the population value. The older SSR did not differ significantly or marginally significantly from the population value for androphilic or gynephilic/biphilic women. For the younger SSR in androphilic women, the range was 101:100-102:100 and was marginally significantly different in one case in the opposite direction of marginally fewer younger brothers than expected. There was no significant difference in SSR, older SSR, or younger SSR between androphilic women and gynephilic/biphilic women (Table 5). Thus, androphilic women tended to have less brothers compared to the population value, but there were no other significant differences in androphilic women or gynephilic/biphilic women, suggesting the SSR was generally comparable to the population value.

### Logistic Regressions: FBO

Results of the logistic regression analyses controlling for age, ethnicity, and education, in men are shown in Table 6. The omnibus was significant in all four analyses. On step 1, modified proportion of older brothers trended toward significance in two analyses, although if a one-tailed test was used, modified proportion of older brothers would be significant in all four analyses. Thus, there was some support that androphilic/biphilic men were more likely to have older brothers than gynephilic men. Modified proportion of younger brothers was not significant in all four analyses on step 1 and would not be significant with a one-tailed test. Modified proportion of older sisters, modified proportion of younger sisters, and modified proportion of same-age sisters were not significant on step 2.

Using Complex Samples, without biphilic men, modified proportion of older brothers and modified proportion of younger brothers were not significant ($p > .800$). Including biphilic men, modified proportion of older brothers was not significant ($p = .144$) and modified proportion of younger brothers was significant ($p = .039$), although modified proportion of older brothers trends toward significance with a one-tailed test. Full results can be obtained from the first author. Thus, results from Complex Samples analyses were somewhat similar to results from analyses not using Complex Samples: androphilic/biphilic men were somewhat more likely to have older brothers than gynephilic men. They were also more likely to have younger brothers.

In women, all four omnibus tests were significant ($\chi^2(5) = 11.62$ to 39.77, $p = .040$ to $<.001$), however modified proportion of older brothers (B = −0.02 to 0.47, Wald = 0.004 to 0.85, $p = .357$ to .950, OR = 0.86 to 1.60) and modified proportion of younger brothers (B = −0.02 to 0.23, Wald = 0.003 to 0.24, $p = .627$ to .955, OR = 0.98 to 1.26) were not

significant in all tests. In two analyses, on step 2, modified proportions of same-age sisters (B =1.30 to 1.35, Wald = 6.75 to 7.22, $p$ = .007 to .009, OR = 3.67 to 3.85) was significant such that gynephilic/biphilic women were more likely to have same-age sisters. These results, and the results from the OBOR and SSR analyses, suggest no older brother effect within women and thus Complex Samples analyses were not conducted.

### Logistic Regressions: Only-Child Status

The numbers and percent of only-children by sexual orientation are found in Table 1 for men and Table 2 for women. In gynephilic men, about 25% were only-children and in androphilic/biphilic men, 22-25% were only-children. In androphilic women, about 23% were only-children, whereas in gynephilic/biphilic women, about 28-30% were only-children.

In men, all four omnibus tests were significant ($\chi^2(4)$ = 18.62 to 30.22, $p$ = .001 to <.001), however the only-child variable (B = −0.10 to 0.06, Wald = 0.04 to 0.31, $p$ = .576 to .836, OR = 0.90 to 1.06) was not significant in all tests. These results suggest no only-child effect within men and thus Complex Samples analyses were not conducted.

In women, all four omnibus tests were significant (see Table 7). Only-child status was significant in two analyses and marginally significant in the remaining two analyses, suggesting that gynephilic/biphilic women are more likely to be only-children than androphilic women, especially when including biphilic women.

Using Complex Samples, without biphilic women, only-child status was not significant ($p$ = .229), but including biphilic women, only-child status was significant ($p$ = .039). Full results can be obtained from the first author. Thus, results from Complex Samples analyses were somewhat similar to results from analyses not using Complex Samples: gynephilic/biphilic women were somewhat more likely to be only-children than androphilic women.

## DISCUSSION

Regarding the FBO effect, we predicted replication of the effect in men, but not in women in the Add Health data. In women, older brothers were not related to sexual orientation in OBOR and logistic regression analyses. In men, results from the OBOR and logistic regressions provided some support for this hypothesis. In men, the OBOR ranged from 1.19 to 1.29 across the samples. Blanchard (2018a) found an average OBOR of 1.27 for non-feminine/cisgender groups and our range contains this same OBOR. Also, Blanchard (2018c) showed that incorporating our results from Subsample I into a meta-analysis of probability samples contributed to a significant OBOR of 1.21. Thus, the OBOR in absolute number is as expected. Also, using a one-tailed approach, some of the results were significant. Given the results of Blanchard (2018a) and a large body of evidence supporting the FBO effect in men, use of a one-tailed approach is justified, although the effects were not fully supportive and two-tailed tests did not achieve significance.

In the logistic regression analyses, modified proportion of older brothers trended toward significance in two analyses, but with a one-tailed test modified proportion of older brothers

would be significant in all analyses. Also, modified proportion of younger brothers, modified proportion of older sisters, modified proportion of younger sisters, and modified proportion of same-age sisters were not significant predictors. Complex Samples generally replicated these results, but with the addition of biphilic men only, and modified proportion of younger brothers was also significant. Thus, there was some support that androphilic/biphilic men were more likely to have older brothers than gynephilic men.

Regarding SSR, we predicted that androphilic men will have a greater SSR than gynephilic men and it will be driven by an older SSR, whereas in women, we predicted a lower SSR in gynephilic women compared to androphilic women based on VanderLaan et al. (2014). Results with the SSR were supportive of our prediction in men: androphilic/biphilic men tended to have a higher SSR, which was driven by the older SSR, compared to the population value and compared to gynephilic men. The SSR ranged from 134:100-138:100 in androphilic/biphilic men, whereas in gynephilic men it was 108:100, which is close to the population value of 106:100. The older SSR for androphilic/biphilic men were high and ranged from 174:100-197:100 whereas the older SSR for gynephilic men ranged from 115:100-116:100. These SSR results are supportive of the FBO effect in men; however, it is somewhat difficult to place them in context with other SSR results given a meta-analysis has not been conducted for SSR. Blanchard (1997) found that the SSR of male-to-female androphilic transsexuals was 134:100, but for more masculine androphilic men, SSR was 108:100. VanderLaan et al. (2014) found that androphilic boys who experience gender dysphoria had a SSR of 128:100. Thus, our results seem to be similar to those of previous research where the men were more feminine, suggesting that perhaps the sample of androphilic/biphilic men in the Add Health data may be more feminine, on average. If so, we would expect the OBOR to be in the range of the feminine samples included in Blanchard (2018a; i.e., close to 1.85), which was not the case. Nevertheless, the SSR results presented here were supportive of the FBO effect in men, and thus demonstrate the value of including this biodemographic variable when examining sexual orientation.

In women, however, the SSR results were not supportive of our prediction: androphilic women tended to have less brothers compared to the population value, otherwise the SSR was generally comparable to the population value in androphilic and gynephilic/biphilic women. VanderLaan et al. (2014) found a SSR of 81:100 in gynephilic girls who experience gender dysphoria, whereas we found a SSR of 94:100-96:100 in androphilic women. In gynephilic/biphilic women, the SSR ranged from 95:100-105:100, but was not significantly different from the population value. It is possible that the SSR results in gynephilic/biphilic women may not have reached significance due to a discrepancy in sample size between gynephilic/biphilic and androphilic women (see Table 2 and see discussion below about sample size in the Add Health data). Nevertheless, our results did not produce a SSR that was as low as VanderLaan et al. (2014) and, along with the results from the OBOR and logistic regressions, does not provide support for an FBO effect in women.

Regarding only-child status, we predicted that gynephilic women and androphilic men would be more likely to be only-children compared to androphilic women and gynephilic men, respectively (Blanchard, 2012a; Skorska et al., 2017; VanderLaan et al., 2014; VanderLaan et al., 2015). There was no only-child effect within men, but within women,

gynephilic/biphilic women were more likely to be only-children than androphilic women, particularly when including biphilic women. The results within women held using Complex Samples when biphilic women were included in analyses. Our hypothesis was thus largely supported within women, but not supported in men. Thus, an only-child effect, but not an FBO effect, was found in women.

Sample size likely plays a role in some of the non-significant and marginally significant findings. There is a large discrepancy between the size of the gynephilic men and the androphilic men groups (see Table 1) and between the size of the androphilic women and gynephilic women groups (see Table 2). Further, the results from Subsample I, which contained a larger sample, especially of androphilic men and gynephilic women, were generally more supportive of each of the hypotheses. In addition, when androphilic and biphilic men were combined, the effects were generally more supportive as well. A post-hoc power analysis conducted using G*Power version 3.1.9.3 (Faul, Erdfelder, Lang, & Buchner, 2007) for an Exact test of proportions for two independent groups (two-tailed, $\alpha = .05$, odds ratio = 1.29, expected proportion of .50, $n_{gynephilic\ men} = 6562$, $n_{androphilic\ men} = 174$) produced power of 0.38. Including the biphilic men in the above post-hoc power analysis brought the power to .47. For a one-tailed test, post-hoc power is .51 and .59, respectively. Last, the significant results of Blanchard (2018c) that include our Subsample I results suggest that with less discrepancy between groups and a generally larger sample size after pooling the results from several studies, our results contribute to the significant finding in Blanchard (2018c). The FBO effect is small in magnitude (Blanchard et al., 1998)—possibly because numerous other factors also influence sexual orientation—and thus a large sample is required to provide sufficient power to find it. A large sample of gynephilic men is helpful, but not having an equivalently large sample of androphilic men is a limitation. Given the patterns of results across all analyses, logistic regressions for FBO and only-child status and SSR results are also likely impacted by the discrepancy in sample size between the two focal groups.

It is somewhat puzzling that the OBOR and logistic regression results were less supportive than the SSR results in men. One issue may be the sample size discrepancy between the groups of men described previously; however, one would expect that the SSR results would be less supportive if sample size is the only issue. SSR of course is calculated differently than OBOR. For example, there may be fewer older sisters in androphilic men, which elevated the older SSR more than in gynephilic men. Another problem may be the recruitment of participants in the Add Health study. It is possible that the Add Health data may not be representative, at least in terms of sexual orientation, or in terms of family structure. The older SSRs were high in both gynephilic and androphilic/biphilic men. Furthermore, in the 1990s it was estimated that about 20% of individuals from the United States do not have at least one sibling (Lavin, 1991; see also Olson, 2015), whereas we found that about 25% of gynephilic men, 22-25% of androphilic/biphilic men, 23% of androphilic women, and 28-30% of gynephilic/biphilic women were only-children. Thus, there is a possibility that the Add Health data contains a lower family size than expected, making the FBO effect more difficult to detect. An additional challenge with the Add Health data is that the sibling variables utilized were estimated from questions about who was living with the participant at the time, instead of estimates of the numbers of each category of

siblings. It is possible that participants had additional biological siblings that are not living with them. It is also possible that families expanded beyond what was reported in the household roster in 1995. For example, additional siblings may have been born after 1995 that would not be reflected in the numbers reported here. In sum, it is important to keep in mind that the Add Health data, even when corrections and sampling weights are applied, likely have limitations, particularly for sibling and sexual orientation variables.

In the current study, we found modest evidence for an FBO effect in men with the OBOR, logistic regression, and SSR results, but no evidence of an only-child effect, supporting existing evidence for the FBO effect in men (e.g., Blanchard & Bogaert, 1996; Blanchard 2018a; 2018c). Although the results of Francis (2008) have sometimes been used to claim there is no support for the FBO effect using the Add Health data (e.g., Blanchard, 2014; VanderLaan et al., 2014; Zietsch, 2018), Francis (2008) did indicate that multiple older brothers would likely be significantly associated with an androphilic sexual orientation at $p < .10$. Here, we have shown modest support for the FBO effect in this data set. Given the limitations with the Add Health data, and a small effect size for the FBO effect, it is not surprising that modest support was found.

The hypothesized explanation for the FBO effect in men is the MIH, which has recently received direct empirical support through the study conducted by Bogaert and colleagues (2018): controlling for total numbers of pregnancies, mothers of androphilic men with older brothers had the highest anti-NLGN4Y concentrations compared to mothers of androphilic men without older brothers, mothers of gynephilic men, and women with no sons. It is intriguing that only-child status was not related to sexual orientation in men given the results of previous studies (Blanchard, 2012a; Skorska et al., 2017; VanderLaan et al., 2015). Interestingly, these studies also focused on the purported mechanism for the only-child findings by also investigating birth weight alongside only-child status. Some research has indicated that the FBO effect may only apply to a subset of androphilic men (Swift-Gallant et al., 2018; Wampold, 2018) and other mechanisms, such as prenatal androgens, may apply to other subset(s) of androphilic men (e.g., Skorska & Bogaert, 2017a; Swift-Gallant, Coome, Skorska, Monks, & VanderLaan, 2018; but see Breedlove, 2017). Perhaps in those androphilic men who are only-children, the immune mechanism associated with a lower birth weight (see below) may be the mechanism behind their only-child status, but, in addition to only affecting a subset of androphilic men, this mechanism may not be as common in androphilic men as it is in gynephilic women.

Indeed, in the current study, we found evidence for an only-child effect in women, but no evidence for an FBO effect. Gynephilic/biphilic women were more likely to be only-children than androphilic women. This finding supports Schagen et al. (2012) and VanderLaan et al. (2014), who found that adolescent girls who experience gender dysphoria were more likely to be only-children. A hypothesized explanation for the only-child effect is an immunological effect independent of the FBO effect (Blanchard, 2012b). A maternal immune response to a fetus was associated with a reduced weight in the newborn and with increased miscarriages in later pregnancies (Kahn & Baltimore, 2010; Nielson, 2011). Support for this independent immune response was found in Blanchard (2012a) by showing that seven gynephilic women and six androphilic men who were only-children had a lower

birth weight than androphilic women and gynephilic men who were only-children and all participants who were the oldest child (cf. VanderLaan et al., 2015). As mentioned previously, in the smaller subset of androphilic men who are only-children, the potential immune mechanism associated with a lower birth weight may be the mechanism behind their only-child status, but this may not be as common in androphilic men as it is in gynephilic women, supporting the finding that gynephilic women were likely to be only-children, but androphilic men were not. The hypothesized protein in men underlying the only-child immune response in the mother is SMCY (Skorska et al., 2017), which differs from the protein related to the FBO effect (NLGN4Y; Bogaert et al., 2018). A maternal immune response to SMCY may be particularly strong and result in delivery of a limited number of offspring whose sexual orientation may be impacted as well. Given SMCY is a male-specific protein any resultant pregnancies may be more likely to be female rather than male and thus SMCY may also be partially involved in the mechanism underlying the only-child findings in gynephilic women. On the other hand, a maternal immune response may be unlikely to be the only mechanism underlying the development of gynephilic women given SMCY is a male-specific protein. Thus, a maternal immune response directed toward SMCY in previous male fetuses may be unlikely to affect brain development of subsequent female fetuses (which do not contain SMCY or other male proteins).

Apart from the limitations noted above regarding sample size discrepancies and limitations with the Add Health data, other limitations are noteworthy. Causal direction cannot be inferred with the study design used in the current study; however, ethically, studies incorporating experimental designs involving sexual orientation are unlikely to be conducted. Also, we did not directly test any of the purported immune-based (or other) mechanisms and thus other interpretations are possible. We were unable to use Complex Samples for all analyses, although we did attempt to ameliorate this concern by presenting analyses that did not use Complex Samples and analyses that did use Complex Samples where possible. Last, in some analyses individuals who are biphilic were grouped with other sexual orientations. Theoretically, it might be argued that immune mechanisms underlying biphilia would be similar to immune mechanisms underlying androphilia in men and gynephilia in women, given all these groups experience some degree of same-sex attraction. Thus, grouping of biphilic individuals with androphilic men and gynephilic women may be warranted. On the other hand, this is an assumption that can be (and should be) empirically tested and thus, future research should recruit large numbers of individuals who are biphilic to compare to androphilic men, gynephilic women, gynephilic men, and androphilic women. For example, there may be a dosage effect where individuals who are biphilic fall in between androphilic men and gynephilic men (and in between gynephilic women and androphilic women) on the family characteristics presented in the current study and other biomarker research conducted on sexual orientation to date (e.g., height, facial structure, 2D:4D finger ratios; Skorska & Bogaert, 2017a; Skorska, Geniole, Vrysen, McCormick, & Bogaert, 2015; Grimbos, Dawood, Burriss, Zucker, & Puts, 2010). Of course, it is possible that a different mechanism or multiple mechanisms may underlie biphilia, and these mechanism(s) could differ between men and women (e.g., Bailey et al., 2016; Diamond & Alley, 2018). Indeed, men tend to be less biphilic than women (Bailey et al., 2016), at least some biphilic men have biphilic sexual arousal patterns (Rosenthal, Sylva, Safron, & Bailey, 2012), and there is

some evidence that biphilic women have greater rates of adversity in early life than gynephilic women (Diamond & Alley, 2018).

Future research that also should be conducted involves studies utilizing large samples without large discrepancies in group sizes. Although national probability samples are preferable in some ways due to the generalizations that can be made with their sampling strategies, given the relatively small percentage of sexual minorities in the population, there will inevitably be group size discrepancies. Larger national probability samples, national probability samples that oversample sexual minorities, or other non-convenience samples that are large in size (e.g., randomly selecting individuals who are sexual minorities to participate in research studies instead of convenience sampling) and that target the groups of interest would be useful steps forward. Wave V of the Add Health data will be available in the future for analysis and it would be interesting to investigate FBO, SSR, and only-child status using this older population, given androphilic men and gynephilic women may have had more chance to come out with age. Replication of only-child findings in women is necessary. Also, understanding the only-child findings in relation to birth weight in both men and women would be helpful. Last, further research into purported dosage effects for SSR and (potentially) biphilia as it relates to the FBO effect is needed. Understanding the relationship between SSR and the FBO effect and the purported underlying MIH mechanism would significantly move this work forward.

There are several strengths with the current study, including using a national probability sample, conducting analyses in both men and women, controlling for family size in analyses, prioritizing older brothers in analyses, excluding participants who have same-age brothers, not using adoptive and foster siblings in summations of the various sibling variables, using a continuous measure of each of the sibling variables, using Wave IV sexual orientation, conducting some analyses with Complex Samples, and inclusion of SSR to provide some important generalization information about the Add Health data and provide an additional family characteristic related to the FBO effect. Keeping these strengths in mind, in the current study, using the Add Health data we found some partial support for the FBO effect related to men's sexual orientation, such that androphilic/biphilic men were more likely to have older brothers than gynephilic men. We also found that gynephilic/biphilic women were more likely to be only-children than androphilic women. We discussed the challenges of investigating the FBO effect using the Add Health data and provided some useful suggestions for future research, including testing of the purported mechanisms underlying the FBO effect and the only-child effect.

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**Table 1.**

Numbers of each sibling category in men by sexual orientation and subsample.

| **Subsample** | ***n* = 19,820** | | | ***n* = 14,172** | | |
|---|---|---|---|---|---|---|
| **Sexual orientation** | **Gynephilic men** | **Androphilic men** | **Androphilic/ biphilic men** | **Gynephilic men** | **Androphilic men** | **Androphilic/ biphilic men** |
| *n* | 6562 | 174 | 225 | 6228 | 167 | 212 |
| *n*(%) only-children | 1618(24.66) | 39(22.41) | 57(25.33) | 1534(24.63) | 39(23.35) | 53(25.00) |
| Full older brothers | 1553 | 49 | 59 | 1457 | 46 | 56 |
| Full younger brothers | 2289 | 71 | 86 | 2193 | 69 | 83 |
| Full older sisters | 1321 | 26 | 31 | 1269 | 26 | 31 |
| Full younger sisters | 2092 | 48 | 61 | 2014 | 45 | 57 |
| Half older brothers | 169 | 8 | 9 | 143 | 4 | 5 |
| Half younger brothers | 497 | 9 | 12 | 460 | 9 | 12 |
| Half older sisters | 159 | 3 | 5 | 128 | 2 | 4 |
| Half younger sisters | 442 | 15 | 18 | 412 | 15 | 18 |
| Full same-age sisters | 149 | 5 | 7 | 129 | 5 | 6 |
| Half same-age sisters | 4 | 1 | 1 | 4 | 0 | 0 |
| Total older brothers | 1722 | 57 | 68 | 1600 | 50 | 61 |
| Total other siblings | 6953 | 178 | 221 | 6609 | 171 | 211 |

*Note*. Sibling counts represent the sums of siblings across all participants. Total older brothers = full older brothers + half older brothers; Total other siblings = full younger brothers + full older sisters + full younger sisters + half younger brothers + half older sisters + half younger sisters + full same-age sisters + half same-age sisters.

**Table 2.**

Numbers of each sibling category in women by sexual orientation and subsample.

| **Subsample** | | | ***n* = 19,820** | | | ***n* = 14,172** |
|---|---|---|---|---|---|---|
| **Sexual orientation** | **Androphilic women** | **Gynephilic women** | **Gynephilic/biphilic women** | **Androphilic women** | **Gynephilic women** | **Gynephilic/biphilic women** |
| *n* | 7727 | 143 | 334 | 7297 | 136 | 317 |
| *n*(%) only-children | 1803(23.33) | 42(29.37) | 100(29.94) | 1694(23.22) | 39(28.68) | 94(29.65) |
| Full older brothers | 1710 | 29 | 59 | 1643 | 29 | 59 |
| Full younger brothers | 2572 | 42 | 94 | 2467 | 42 | 94 |
| Full older sisters | 1647 | 28 | 57 | 1549 | 27 | 53 |
| Full younger sisters | 2546 | 33 | 84 | 2408 | 32 | 81 |
| Half older brothers | 213 | 5 | 14 | 182 | 5 | 14 |
| Half younger brothers | 640 | 10 | 33 | 577 | 10 | 32 |
| Half older sisters | 213 | 7 | 15 | 183 | 6 | 13 |
| Half younger sisters | 626 | 10 | 33 | 577 | 10 | 29 |
| Full same-age sisters | 438 | 8 | 22 | 357 | 7 | 19 |
| Half same-age sisters | 2 | 0 | 0 | 2 | 0 | 0 |
| Total older brothers | 1923 | 34 | 73 | 1825 | 34 | 73 |
| Total other siblings | 8684 | 138 | 338 | 8120 | 134 | 321 |

*Note*. Sibling counts represent the sums of siblings across all participants. Total older brothers = full older brothers + half older brothers; Total other siblings = full younger brothers + full older sisters + full younger sisters + half younger brothers + half older sisters + half younger sisters + full same-age sisters + half same-age sisters.

**Table 3.**

Older brother odds ratio (OBOR) results in men and in women.

| | | | | |
|---|---|---|---|---|
| | | **Men** | | |
| **Subsample** | ***n* = 19,820** | | ***n* = 14,172** | |
| Comparison | Gynephilic men vs. androphilic men | Gynephilic men vs. androphilic/biphilic men | Gynephilic men vs. androphilic men | Gynephilic men vs. androphilic/biphilic men |
| OBOR (men) | 1.29 | 1.24 | 1.21 | 1.19 |
| 95% CI | 0.96-1.75 | 0.94-1.64 | 0.88-1.66 | 0.89-1.60 |
| 90% CI | 1.00-1.67 | 0.98-1.57 | 0.92-1.58 | 0.94-1.52 |
| *p* | *.096* | .124 | .247 | .231 |
| | | Women | | |
| Comparison | Androphilic women vs. gynephilic women | Androphilic women vs. gynephilic/biphilic women | Androphilic women vs. gynephilic women | Androphilic women vs. gynephilic/biphilic women |
| OBOR (women) | 1.11 | 0.98 | 1.13 | 1.01 |
| 95% CI | 0.76-1.62 | 0.75-1.26 | 0.77-1.65 | 0.78-1.31 |
| 90% CI | 0.81-1.53 | 0.79-1.21 | 0.82-1.55 | 0.81-1.26 |
| *p* | .581 | .849 | .531 | .929 |

*Note*. OBOR (men) = (androphilic men's number of older brothers ÷ androphilic men's number of other siblings) ÷ (gynephilic men's number of older brothers ÷ gynephilic men's number of other siblings); OBOR (women) = (gynephilic women's number of older brothers ÷ gynephilic women's number of other siblings) ÷ (androphilic women's number of older brothers ÷ androphilic women's number of other siblings). CI = confidence interval. **Bold**, $p < .05$, *italics*, $p < .10$, *p*-values are two-tailed and for $\alpha = .05$ (i.e., the 95% CI).

**Table 4.**

Sibling sex ratio (SSR), including comparison to population value, in men and in women.

| **Men** | | | | | | |
|---|---|---|---|---|---|---|
| **Subsample** | | ***n* = 19,820** | | | ***n* = 14,172** | |
| Sexual orientation | Gynephilic men | Androphilic men | Androphilic/ biphilic men | Gynephilic men | Androphilic men | Androphilic/ biphilic men |
| *n* | 6562 | 174 | 225 | 6228 | 167 | 212 |
| Sibling sex ratio | 108:100 | 140:100 | 135:100 | 108:100 | 138:100 | 134:100 |
| *z*-ratio | 0.93 | 2.03 | 1.98 | 0.62 | 1.85 | 1.88 |
| *p* | .35 | **.04** | **.05** | .54 | *.06* | *.06* |
| Older sibling sex ratio | 116:100 | 197:100 | 189:100 | 115:100 | 179:100 | 174:100 |
| *z*-ratio | 2.61 | 2.64 | 2.74 | 2.09 | 2.12 | 2.27 |
| *p* | **.009** | **.008** | **.006** | **.04** | **.03** | **.02** |
| Younger sibling sex ratio | 110:100 | 127:100 | 124:100 | 109:100 | 130:100 | 127:100 |
| *z*-ratio | 1.31 | 0.99 | 0.96 | 1.09 | 1.1 | 1.08 |
| *p* | .19 | .32 | .34 | .28 | .14 | .28 |
| Women | | | | | | |
| Sexual orientation | Androphilic women | Gynephilic women | Gynephilic/ biphilic women | Androphilic women | Gynephilic women | Gynephilic/ biphilic women |
| *n* | 7727 | 143 | 334 | 7297 | 136 | 317 |
| Sibling sex ratio | 94:100 | 100:100 | 95:100 | 96:100 | 105:100 | 102:100 |
| *z*-ratio | −6.27 | −0.31 | −1.09 | −4.98 | 0.01 | −0.33 |
| *p* | <**.001** | .76 | .28 | <**.001** | .99 | .74 |
| Older sibling sex ratio | 103:100 | 97:100 | 101:100 | 105:100 | 103:100 | 111:100 |
| *z*-ratio | −0.76 | −0.24 | −0.19 | −0.17 | 0.01 | 0.16 |
| *p* | .45 | .81 | .85 | .87 | .99 | .87 |
| Younger sibling sex ratio | 101:100 | 121:100 | 109:100 | 102:100 | 124:100 | 115:100 |
| *z*-ratio | −1.82 | 0.54 | 0.12 | −1.5 | 0.65 | 0.53 |
| *p* | *.07* | .59 | .90 | .13 | .52 | .60 |

*Note.* SSR = brothers/sisters*100; Older SSR = older brothers/older sisters*100; Younger SSR = younger brothers/younger sisters*100. **Bold**, $p < .05$, *italics*, $p < .10$, *p*-values are two-tailed.

**Table 5.**

Comparisons of sibling sex ratio (SSR) by sexual orientation in men and in women.

| Men | | | | |
|---|---|---|---|---|
| **Subsample** | ***n* = 19,820** | | ***n* = 14,172** | |
| Comparison | Gynephilic men vs. androphilic men | Gynephilic men vs. androphilic/biphilic men | Gynephilic men vs. androphilic men | Gynephilic men vs. androphilic/biphilic men |
| SSR *z* | −1.92 | −1.83 | −1.79 | −1.80 |
| SSR *p* | *.06* | *.07* | *.07* | *.07* |
| Older SSR *z* | −2.30 | −2.34 | −1.90 | −2.00 |
| Older SSR *p* | **.02** | **.02** | *.06* | **.05** |
| Younger SSR *z* | −0.85 | −0.79 | −1.00 | −0.94 |
| Younger SSR *p* | .40 | .43 | .32 | .35 |
| Women | | | | |
| Comparison | Androphilic women vs. gynephilic women | Androphilic women vs. gynephilic/biphilic women | Androphilic women vs. gynephilic women | Androphilic women vs. gynephilic/biphilic women |
| SSR *z* | −0.41 | −0.10 | −0.57 | −0.60 |
| SSR *p* | .68 | .92 | .57 | .55 |
| Older SSR *z* | 0.26 | 0.12 | 0.09 | −0.28 |
| Older SSR *p* | .80 | .91 | .93 | .78 |
| Younger SSR *z* | 0.64 | −0.25 | 0.52 | −0.46 |
| Younger SSR *p* | .52 | .80 | .60 | .65 |

*Note*. **Bold**, $p < .05$, *italics*, $p < .10$, *p*-values are two-tailed.

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**Table 6.**

Logistic regressions with modified proportion of older brothers (MPOB) and other sibling variables controlling for age, education and ethnicity in men.

| **Subsample** | ***n* = 19,820** | | | | | | | | ***n* = 14,172** | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Comparison** | **Gynephilic men vs. androphilic men** | | | | **Gynephilic men vs. androphilic/biphilic men** | | | | **Gynephilic men vs. androphilic men** | | | | **Gynephilic men vs. androphilic/biphilic men** | | | |
| Om $\chi^2$(df) | 32.87(5) | | | | 22.58(5) | | | | 33.99(5) | | | | 22.34(5) | | | |
| Om *p* | **<.001** | | | | **<.001** | | | | **<.001** | | | | **<.001** | | | |
| Nag $R^2$ | .023 | | | | .013 | | | | .025 | | | | .014 | | | |
| | B | W | *p* | OR | B | W | *p* | OR | B | W | *p* | OR | B | W | *p* | OR |
| S1: MPOB | 0.88 | 3.65 | *.056* | 2.42 | 0.74 | 3.27 | *.071* | 2.10 | 0.82 | 2.88 | *.090* | 2.26 | 0.74 | 3.01 | *.083* | 2.10 |
| S1: MPYB | 0.67 | 2.52 | .112 | 1.95 | 0.54 | 2.13 | .145 | 1.72 | 0.70 | 2.69 | .101 | 2.01 | 0.58 | 2.35 | .126 | 1.79 |
| S1: Age | −0.01 | 0.11 | .739 | 0.99 | −0.03 | 0.45 | .504 | 0.98 | 0.01 | 0.03 | .867 | 1.01 | −0.01 | 0.03 | .858 | 0.99 |
| S1: Eth | 0.17 | 1.17 | .280 | 1.18 | 0.17 | 1.44 | .231 | 1.18 | 0.11 | 0.46 | .497 | 1.12 | 0.11 | 0.64 | .424 | 1.12 |
| S1: Edu | 0.91 | 23.31 | **<.001** | 2.49 | 0.60 | 15.39 | **<.001** | 1.83 | 0.99 | 24.92 | **<.001** | 2.70 | 0.65 | 16.28 | **<.001** | 1.92 |
| S2 | *ns* | | | | *ns* | | | | *ns* | | | | *ns* | | | |

*Note*. Om = omnibus, Nag = Nagelkerke, S1 = step 1, MPYB = modified proportion of younger brothers, Eth = ethnicity, Edu = education, S2 = step 2, W = Wald statistic, OR = odds ratio. Gynephilic men were coded 0 and androphilic/biphilic men were coded 1; No college/university education was coded 0 and some college/university education was coded 1; White ethnicity was coded 0 and other ethnicities were coded 1. Complex Samples were not used in these analyses. **Bold**, $p < .05$, *italics*, $p < .10$, *p*-values are two-tailed.

**Table 7.**

Logistic regressions with only-child (OC) status controlling for age, education and ethnicity in women.

| **Subsample** | ***n* = 19,820** | | | | | | | | ***n* = 14,172** | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Comparison** | **Androphilic women vs. gynephilic women** | | | | **Androphilic women vs. gynephilic/biphilic women** | | | | **Androphilic women vs. gynephilic women** | | | | **Androphilic women vs. gynephilic/biphilic women** | | | |
| Om $\chi^2$(df) | 14.48(4) | | | | 44.25(4) | | | | 16.20(4) | | | | 48.31(4) | | | |
| Om *p* | **.006** | | | | **<.001** | | | | **.003** | | | | **<.001** | | | |
| Nag $R^2$ | .011 | | | | .019 | | | | .013 | | | | .022 | | | |
| | B | W | *p* | OR | B | W | *p* | OR | B | W | *p* | OR | B | W | *p* | OR |
| S1: OC | 0.36 | 3.61 | *.058* | 1.43 | 0.38 | 9.45 | **.002** | 1.46 | 0.34 | 3.12 | *.077* | 1.41 | 0.39 | 9.18 | **.002** | 1.47 |
| S1: Age | −0.08 | 2.84 | *.092* | 0.92 | −0.12 | 13.33 | **<.001** | 0.89 | −0.10 | 3.81 | *.051* | 0.91 | −0.13 | 16.19 | **<.001** | 0.88 |
| S1: Eth | 0.41 | 5.87 | **.015** | 1.51 | 0.30 | 6.87 | **.009** | 1.34 | 0.45 | 6.83 | **.009** | 1.57 | 0.33 | 7.96 | **.005** | 1.39 |
| S1: Edu | −0.33 | 3.48 | *.062* | 0.72 | −0.50 | 18.96 | **<.001** | 0.61 | −0.37 | 4.20 | **.040** | 0.69 | −0.53 | 20.54 | **<.001** | 0.59 |

*Note*. Om = omnibus, Nag = Nagelkerke, S1 = step 1, Eth = ethnicity, Edu = education, W = Wald statistic, OR = odds ratio. Androphilic women were coded 0 and gynephilic/biphilic women were coded 1; Not an only-child was coded 0 and only-child was coded 1. No college/university education was coded 0 and some college/university education was coded 1; White ethnicity was coded 0 and other ethnicities were coded 1. Complex Samples were not used in these analyses. **Bold**, $p < .05$, *italics*, $p < .10$, *p*-values are two-tailed.