Published in final edited form as:
J Marriage Fam. 2012 February 1; 74(1): 201–214. doi:10.1111/j.1741-3737.2011.00875.x.

Mate Availability and Women's Sexual Experiences in China

Katherine Trent and **Scott J. South**
University at Albany, State University of New York

Abstract

Data from the 1999–2000 Chinese Health and Family Life Survey were merged with community-level data from the 1982, 1990, and 2000 Chinese censuses to examine the relationship between the local sex ratio (number of men per 100 women) and sexual outcomes among women (N = 1,369). Consistent with hypotheses derived from demographic-opportunity theory, multilevel logistic regression analyses showed that women are more likely to be sexually active, to have had premarital sexual intercourse, to have been forced to have sex, and to test positive for a sexually transmitted infection when there is a relative abundance of age-matched men in their local community. Education, birth cohort, and geographic location also emerged as significant predictors of women's sexual experiences.

Keywords

China; demography; sex ratio; sexual behavior; sexually transmitted diseases

The relative numbers of young men and women in the People's Republic of China have undergone remarkable change over recent decades. China's one-child policy, precipitous declines in fertility, a cultural preference for sons, and the increasing prevalence of sex-selective abortion technology have converged to generate a surplus of men (Banister, 2004; Goodkind, 2004). Many scholars have pointed to an unusually high sex ratio at birth (number of male per 100 female births) in China (e.g., Cai & Lavely, 2003; Coale & Banister, 1994; Gu & Roy, 1995; Hull, 1990; Johansson & Nygren, 1991; Lavely, 2001; Murphy, 2003; Peng & Huang, 1999; Secondi, 2002; Yi et al., 1993; Yuan & Tu, 2004). A normal range of the sex ratio at birth is considered to be between 103 and 107. In 1982, China's sex ratio at birth—between 107 and 108—was already at the high end of this range, and it has increased dramatically since that time. By 1990, the sex ratio at birth had grown to 111.3; by 2001, it was 118 (Poston & Glover, 2005); by 2005, it had reached 120.5 (Li, 2007).

Given these changes in the sex ratio at birth, China is expected to experience a pronounced surplus of adult men relative to adult women as these birth cohorts age (Tuljapurkar, Li, & Feldman, 1995). Some observers have suggested that this increasing population of excess men will have far-reaching social and demographic consequences. For example, Poston and Glover (2005) suggested that the deficit of women will increase commercial sexual activity among men and thus hasten the spread of HIV and AIDS throughout the population (see also Ebenstein & Jennings, 2009; Tucker et al., 2005).

But little systematic empirical research exists on how imbalanced sex ratios influence sexual behaviors in China or in other populations. In this article, we examine the effect of imbalanced sex ratios on several aspects of Chinese women's sexual experiences. Given

Department of Sociology, University at Albany, State University of New York, Albany, NY 12222 (k.trent@albany.edu).

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recent changes and the substantial intercommunity variation in its sex ratio, China represents an opportune case for examining the impact of imbalanced sex ratios on women's sexual outcomes. Our conceptual framework is grounded in *demographic-opportunity theory* (South, Trent, & Shen, 2001; Uecker & Regnerus, 2010), which broadly suggests that a surplus of men will shape both the frequency and form of women's sexual encounters. We tested hypotheses derived from this theory using individual-level data from the Chinese Health and Family Life Survey (CHFLS) merged with community-level data taken from three Chinese censuses. From the Chinese censuses, we created cohort-specific and community-specific sex ratios describing the number of men available to women, and we then attached these sex ratios to the individual records of the female respondents to the CHFLS. We then estimated multilevel logistic regression models linking four outcomes—whether women (a) are sexually active, (b) have been victims of forced sex, (c) have had premarital sex, and (d) test positive for a sexually transmitted infection (STI)—to the age-specific sex ratio in their local community.

Theoretical Background and Hypotheses

A common theoretical framework from which to address the effect of imbalanced sex ratios on sexual and familial behavior is demographic-opportunity theory (South et al., 2001; Uecker & Regnerus, 2010), which considers the distribution of the population by gender, as well as by other critical sociodemographic characteristics (e.g., age and race), to be a defining characteristic of social structure (Blau, 1977). A fundamental premise of demographic-opportunity theory is that the likelihood of social contact between people with different demographic attributes—for example, between women and men—is determined in part by the number of available outgroup members with whom such contacts could occur (South et al., 2001; Uecker & Regnerus, 2010). Thus, demographic-opportunity theory emphasizes how the sheer number of men available to women shapes the frequency and form of women's sexual encounters. When applied to women's sexual experiences, demographic-opportunity theory suggests that the probability of being sexually active, engaging in premarital sexual intercourse, of being the victim of forced sex, and of testing positive for an STI increases along with the number of men in the local population.

Previous tests of demographic-opportunity theory have focused primarily on how imbalanced sex ratios affect marital and familial behavior in the United States. Women's marriage rates are higher in geographic areas that contain more eligible and economically attractive men (e.g., Fossett & Kiecolt, 1993; Lichter, McLaughlin, Kephart, & Landry, 1992; McLaughlin, Lichter, & Johnston, 1993). A surplus of men has also been linked to young women's chances of "marrying up" educationally (Lichter, Anderson, & Hayward, 1995). High sex ratios are associated with an increase in women's likelihood of nonmarital childbearing (Billy & Moore, 1992; South & Lloyd, 1992). A numerical surplus of men ostensibly increases young women's risk of unmarried childbearing by increasing the likelihood that they will engage in premarital intercourse, although this association is likely tempered by an accompanying increase in early marriage (South, 1996). Similar findings have been observed in cross-national studies based on large samples of countries (Barber, 2001; South & Trent, 1988). Imbalanced sex ratios have also been linked to marital dissolution and relationship quality. A surplus of women lowers relationship quality among unmarried parents (Harknett, 2008), though not among married persons more generally (Trent & South, 2003). Several studies have suggested that married couples who are exposed to a surplus of either men or women are more likely to divorce, presumably because in these demographic contexts spouses are especially likely to encounter an attractive alternative to their current partner (McKinnish, 2004; South et al., 2001).

The literature is sparser and less consistent regarding the influence of mate availability on sexual behavior. Consistent with demographic-opportunity theory, Billy, Brewster, and Grady (1994) found a positive association in the United States between the county-level sex ratio and young women's likelihood and frequency of engaging in premarital intercourse; however, Brewster (1994) did not observe a significant association between the neighborhood-level sex ratio and the timing of young Black women's transition to first sexual activity. Also, contrary to the predictions of demographic-opportunity theory, Uecker and Regnerus (2010) found that young women were more likely to engage in sex when they attended college with comparatively few men. Browning and Olinger-Wilbon (2003) found that men engaged in more short-term sexual alliances in neighborhoods that contained comparatively few women.

Imbalanced sex ratios have also been linked to the frequency of violence—in particular, sexual violence—against women. Although the literature is not entirely consistent (O'Brien, 1991; Whaley, 2001), several studies have shown that, within the United States, rape victimization rates are higher in areas characterized by a surplus of men and an attendant deficit of women (Blau & Golden, 1986; Messner & Blau, 1987). In addition, the sex ratio is positively associated with female homicide victimization (Avakame, 1999) and male-on-female intimate partner violence (D'Alessio & Stolzenberg, 2010). Presumably, when women are scarce, men lack the ability to form conventional sexual relationships, and thus some resort to violence to satisfy their sexual needs and maintain control over actual or potential mates. More generally, China's increasingly masculine sex ratio has been argued to be a partial cause of its increasing rate of crime (Edlund, Li, Yi, & Zhang, 2007).

Only recently have studies begun to explore the impact of China's sex ratio imbalance on family and sexual behaviors. South and Trent (2010) found that, among men, a relative surfeit of women in the local marriage market increased the likelihood of engaging in premarital intercourse and of testing positive for an STI while reducing the likelihood of having commercial sexual intercourse. Trent and South (2011) found that, among women, the male-to-female sex ratio in the marriage market was positively associated with early marriage, the probability of engaging in intercourse outside of marriage, and the likelihood of having had more than one sexual partner over one's lifetime. In the current analysis, we extended these studies to examine how the relative number of men available to women influenced women's chances of being sexually active, of having been forced to have sex, and of contracting an STI.

Demographic-opportunity theory implies several hypotheses regarding the impact of the local sex ratio on the nature and consequences of women's sexual encounters. First, demographic-opportunity theory predicts that, when faced with a relative surplus of men, women will be more likely to be sexually active. A relative abundance of men increases the likelihood that women will encounter an attractive sexual partner—particularly, though not exclusively, through marriage—thereby increasing their chances of engaging in sexual intercourse. In contrast, when men are scarce, women's opportunities to attract a sexual partner will be more limited, and hence women will be less likely to engage in sexual intercourse. Thus, demographic-opportunity theory predicts a positive association between the local sex ratio and women's chances of being sexually active.

Second, demographic-opportunity theory implies that women will be more likely to be forced to have sexual intercourse when they are exposed to a numerical surplus of men. When women are in short supply, many men in women's pool of eligibles will be unable to find sexual and marital partners through more conventional, socially sanctioned means. Instead, some men will turn to illicit or criminal behavior, such as visiting commercial sex workers or obtaining sex through physical force. Even among married persons, a numerical surplus of men—and an accompanying deficit of women—may spur sexual violence by husbands, both because men will have few sexual opportunities outside of marriage and as a means of preventing women from exploiting the extramarital opportunities available to them (D'Alessio & Stolzenberg, 2010). For all these reasons, demographic-opportunity theory predicts a positive association between the community sex ratio and women's likelihood of being victims of forced sexual intercourse.

A third hypothesis implied by demographic-opportunity theory is that women's likelihood of engaging in premarital sexual intercourse is greater in communities that contain relatively large numbers of men. A numerical surplus of men increases the chances that women will encounter an attractive sexual partner before marrying. To be sure, there may be countervailing forces at work here, because an abundance of men may also increase the chances that women will meet a potential husband early in life, and marrying young limits the amount of time that women are exposed to the risk of having sexual intercourse prior to marriage. Nevertheless, because most Chinese women marry fairly late in life (Sheng, 2005), this offsetting influence is likely to be minimal.

Finally, we extend demographic-opportunity theory to hypothesize that a numerical surplus of men will increase the risk that women will contract an STI. A surplus of men is likely to increase women's chances of contracting an STI through several pathways. First, a male surplus is likely to increase women's chances of being sexually active, as well as the sheer frequency with which women engage in sexual intercourse; in turn, and all else equal, being sexually active increases the risk of contracting an STI. Moreover, a surplus of men likely influences the nature of women's sexual encounters beyond merely having sexual encounters and the frequency of intercourse. As we argued earlier, a surplus of men is likely to increase women's likelihood of engaging in sexual intercourse prior to marriage. In addition, a surfeit of men and a concomitant shortage of women may also increase women's likelihood of being forced to have sexual intercourse and of engaging in sex with a partner who has visited commercial sex workers. All of these are risk factors for contracting an STI, including HIV and AIDS (Gil, Wang, Anderson, Lin, & Wu, 1996; Merli, Hertog, Wang, & Li, 2006; Tucker, Ren, & Sapio, 2010; Xiao, Kristensen, Sun, Lu, & Vermund, 2007). In contrast, when men are relatively scarce, women will have fewer opportunities to engage in sexual intercourse, be less likely to have intercourse forced on them, and have sex with safer partners, thereby diminishing their risk of contracting an STI.

In assessing the impact of imbalanced sex ratios on women's sexual experiences, it is important to consider the broader context of family and sexual behavior in China, in particular as it reflects ideological and cultural changes that have occurred over recent decades. As China has moved toward becoming a less ideologically controlled society, family-related values have become increasingly diverse (Sheng, 2005). These secular changes in values associated with family and sexual behavior have likely influenced women's sexual experiences, even though women's sexual behavior in China remains conservative compared with the United States and many other countries (Parish, Laumann, & Mojola, 2007). Nonetheless, in the currently more liberal environment, Chinese women now have higher status and greater latitude in their sexual behaviors, including the freedom to choose their sex partners (Parish et al., 2007). Thus, the chances of being sexually active, engaging in premarital sex, experiencing forced sex, and of contracting an STI are likely to vary across birth cohorts.

Other individual- and community-level characteristics are also likely to be associated with women's sexual experiences. Higher levels of education may lead to more liberal attitudes that afford women more freedom in their sex lives (Parish et al., 2007; Sheng, 2005). Norms regulating women's sexual behavior are likely to vary across China's geographic regions, being particularly more liberal in the more modernized South and East. We controlled for community-level education and urbanization, because both factors could potentially confound an association between the sex ratio and the outcome variables. Rural communities with low levels of education are perhaps likely to exhibit high levels of son preference (and thus high sex ratios) while adhering to norms that constrain women's sexual freedom.

Method

We tested the hypotheses developed earlier using data from the CHFLS in conjunction with community-level data from three Chinese censuses. The CHFLS is a nationally representative survey (with the exception of Hong Kong and Tibet) of 3,821 Chinese adults ages 20 to 64 (Population Research Center, 2006). The CHFLS was administered between August 1999 and August 2000. Modeled in large part on the U.S. National Health and Social Life Survey (Laumann, Gagnon, Michael, Michaels, 1994), the CHFLS focuses on sexual and family-related behaviors and attitudes (Parish et al., 2003).

For this analysis we selected female CHFLS respondents between the ages of 20 and 44 (N = 1,369). We focused on this age range in part because women older than 44 are not likely to have experienced the numerical surplus of men experienced by younger cohorts. Moreover, our measurement strategy required that we estimate the relative numbers of men "available" to these women when they were age 20. The earliest available China census containing the requisite information is for 1982, and thus it was not possible to estimate with confidence the community-and cohort-specific sex ratio for women who were older than 44 at the date of the CHFLS administration.

Dependent Variables

We examined the impact of the relative number of men available to the female CHFLS respondents on four dimensions of women's sexual encounters and their outcomes. All of these variables are dichotomous. *Sexually active* is a dichotomous variable that we scored 1 if the respondent reported having had sexual intercourse in the past year. *Forced sex* is a dichotomous variable that we scored 1 if the respondent reported ever having been forced to have sex against her will. *Premarital sex* is a dichotomous variable that we scored 1 for respondents who reported having had sexual intercourse prior to marriage. Although sensitive questions about sexual behavior may be prone to misreporting, the CHFLS used procedures to ensure the confidentiality of responses, including having respondents enter responses on a laptop computer. Moreover, nonresponses were quite rare; less than 1% of the CHFLS respondents refused to answer the items used to construct these measures. Finally, the CHFLS respondents were asked to provide a urine sample to be tested for the presence of gonorrhea, chlamydia, and trichomonas infections. Over 90% of the respondents provided a urine sample. The fourth dependent variable (whether the respondent had an STI) is a dichotomous variable that we scored 1 for respondents who tested positive for one of these STIs. A limitation of these measures is that we were unable to specify exactly when during their lives women experienced these sexual events or contracted an STI; we address the likely implications of this limitation in a later section of this article.

Independent Variables

Our focal independent variable was the sex ratio, expressed here as the number of men per 100 women. The relevant pool of men available to serve as sexual partners for women is of course circumscribed both by geography and by age. To circumscribe these pools of eligible mates geographically, we coded the county or county equivalent (e.g., urban district, county-level city) for each of the CHFLS respondents. For county-level cities that were under prefecture-level cities and *shixiaqu*, we used data for the entire prefecture-level city (in essence, a large city or metropolitan area). For county-level cities that were under the province, and for noncity counties, we used data at the county level. These geographic approximations of community correspond to the spatially defined marriage markets (e.g., metropolitan areas, labor market areas, or nonmetropolitan counties) used in much U.S. research on the impact of imbalanced sex ratios (e.g., Lichter et al., 1992). The 1,369 CHFLS respondents in our sample were distributed across 37 such communities, with community-specific sample sizes ranging from 14 to 192.

To circumscribe the relevant pool of men by age, and to take into account the fact that the sexual behaviors that serve as dependent variables could have occurred many years before the administration of the CHFLS, we assigned to each female respondent a 7-year sex ratio with a 2-year staggering of the numerator (number of men) and denominator (number of women) when the respondent was age 20. This 2-year staggering corresponds to the age difference between spouses in China (Porter, 2006). Thus, we defined the sex ratio as the number of men ages 17 to 23 divided by the number of women ages 15 to 21. We used data from the full-count 2000 China census (China Data Center, 2004) and the 1% samples from the 1982 and 1990 censuses (China Population and Information Research Center, 2008), along with standard techniques of interpolation and extrapolation, to estimate the value of this community-specific sex ratio for each female CHFLS respondent when she was age 20. A few communities could not be identified in the 1982 or 1990 censuses. For women in these communities, we substituted the age-specific sex ratio observed in the 2000 census. We also observed some extreme values of the sex ratio, likely a consequence of sampling variability in the smaller communities. To limit the influence of these extreme values, we bottom-coded and top-coded the values of the sex ratio at 80 and 120, respectively.

We included several other explanatory variables in our models. Educational attainment was captured by a series of dummy variables ranging from *never attended school* to *attended university or graduate school*. To capture age-related or historical trends in sexual behavior, the models included dummy variables for decadal birth cohort (1950s, 1960s, and 1970s, with the 1950s serving as the reference category). We included a dummy variable for whether respondents reported residing in an urban area (a county-level city or larger) when they were age 14. A separate dummy variable differentiated residents of the generally more modernized south and east coasts of China from other areas. We controlled for two community-level characteristics that could potentially confound an association between the sex ratio and the outcome variables. We computed both of these controls by aggregating responses from all CHFLS respondents to the community level. *Community education* was measured by the proportion of the CHFLS respondents who attended at least senior high school. *Community percent urban* was the percentage of the community members who resided in an urban locale, defined as a village or neighborhood in which fewer than 15% of the workers are farmers. In Table 1 we present definitions for all the variables used in our analyses.

Analytical Strategy

To examine the effect of the availability of men on women's sexual behavior, we estimated multilevel logistic regression models. These models include random intercepts that take into account the clustering of respondents within the 37 communities (Raudenbush & Bryk, 2002). Allowing the model intercepts to vary randomly across communities allows for a correlation between respondents in the same communities. Multilevel models generate significance tests for the community-level variables that are based on the correct degrees of freedom.

Results

In Table 1 we present (weighted) descriptive statistics for all variables used in the analysis. Approximately 89% of the female CHFLS respondents aged 20 to 44 reported being sexually active. Almost 7% of the respondents reported having been forced to have sex against their will at some point in their life, and approximately 12% of the respondents reported having engaged in sexual intercourse prior to marriage. Of the respondents who agreed to provide a urine sample, fewer than 5% tested positive for a gonorrheal, chlamydial, or trichomoniasis infection. (The majority of the positive test results were for chlamydial infection.)

Descriptive statistics for the primary explanatory variable indicated that, on average, there were approximately 108 men aged 17 to 23 per 100 women aged 15 to 21 in the respondents' communities when these respondents were age 20. Thus, on average these women tended to face a surplus of men in their local community during early adulthood. The highest grade attended by the plurality of respondents was junior high school. More than 10% of respondents had never attended school, and fewer than 5% attended college. Fourteen percent of the respondents were born in the 1950s (and were thus ages 40–44 at the time of the CHFLS administration), 44% were born during the 1960s (and were thus ages 30–39 at the time of the survey), and 42% were born during the 1970s (and were thus ages 20–29 at the time of the survey). Fifteen percent of the respondents resided in an urban area at age 14, and 11% resided in a community on the east or south coast of China at the time of the CHFLS. In the typical community, approximately 30% of the adult population had attended at least senior high school, and 45% resided in a village or neighborhood categorized as urban.

In Table 2 we present a series of logistic regression models relating each of the four dimensions of women's sexual outcomes to the community- and cohort-specific sex ratio and the other explanatory variables. Model 1 shows the results for being currently sexually active. The coefficient for the sex ratio was positive and statistically significant. As predicted by demographic-opportunity theory, the greater the number of men available to women, the greater was the likelihood that those women had recent sexual intercourse. To illustrate the magnitude of the effect, a one-unit difference in the male-to-female sex ratio increases the odds that women are sexually active by 3.9% $\{= [e^{(.038)} - 1] \times 100\}$. Perhaps a more useful metric for assessing the magnitude of this effect is to use recent changes in the sex ratio at birth. Between 1982 and 2005, the sex ratio at birth in China increased by about 12.5 men per 100 women: from 108 to 120.5. A change of this magnitude in the young adult sex ratio would increase the odds that women are sexually active by 61% $(= \{[e^{(.038)(12.5)}] - 1\} \times 100)$.

The coefficients for the dummy variables for birth cohort indicated that the 1960s cohort was significantly (at a borderline level) more likely, and the youngest cohort (1970s) was significantly less likely, than the oldest cohort (1950s) to have had sexual intercourse in the past year. Members of the 1960s cohort are probably more likely than members of the 1950s cohort to be sexually active because the former are younger and less likely to be widowed or divorced. The difference between the youngest and the oldest cohort likely stems from the fact that many members of the youngest cohort had yet to marry and thus faced a lower risk of sexual intercourse. Respondents who resided in an urban area at age 14 were (at a borderline level) significantly less likely than respondents with a rural childhood to be currently sexually active. None of the coefficients for the other independent variables in Model 1 was statistically significant.

Model 2 of Table 2 presents the results for the likelihood that the CHFLS female respondents had ever been forced to have sexual intercourse. The coefficient for the sex ratio was again positive and statistically significant. When women were faced with an abundance of men in their local marriage market at age 20, and men were correspondingly faced with comparatively few women, women were more likely to report having been forced against their will to engage in sexual intercourse. A difference of 1 man per 100 women raises the odds that women have been forced to have sex by 1.8% {= [$e^{(.018)}$ − 1] × 100}. Applying the simulation described earlier, an increase in the sex ratio of 12.5 men per 100 women would increase the odds that women have been forced to have sex by about 25% (= {[$e^{(.018)(12.5)}$] − 1} × 100).

Several of the other independent variables also evinced a significant association with the likelihood of women reporting having been forced to have sex. Women who attended at least junior school high school were significantly more likely than women who never attended school to report having experienced forced sex. The positive effect of education on forced sex is perhaps due to the fact that women with more education are both at higher risk of sexual violence and more likely to perceive and report the same type of sexual encounter as forced. Compared to women with less education, women with more education presumably would have more independence and freedom of movement that may increase their risk of forced sex; they may also be more sensitive to the issue of sexual violence and more empowered to report its occurrence. Members of the youngest (1970s) cohort were significantly less likely than members of the oldest (1950s) cohort to have experienced forced sex, a possible function of the former cohort's shorter duration of exposure to the lifetime risk of sexual violence. Residents of the south or east coasts and (at a borderline level) more urban communities were significantly less likely than residents of other regions and less urban places to report having had forced sex.

Model 3 of Table 2 is a logistic regression model of the odds that women have engaged in premarital sexual intercourse. As we anticipated on the basis of demographic-opportunity theory, the coefficient for the sex ratio was positive and statistically significant. The more men who were "available" to women, the greater the likelihood that those women have engaged in premarital sexual intercourse, despite China's strong cultural proscriptions against sex prior to marriage. The effect of the sex ratio was moderate in strength. A difference of 1 man per 100 women increases the odds that women will have had premarital sexual intercourse by 1.7% {= [$e^{(.017)}$ − 1] × 100}, and an addition of 12.5 men per 100 women increases the odds by one quarter (= {[$e^{(.017)(12.5)}$] − 1} × 100).

The likelihood of having had premarital intercourse increased monotonically across birth cohorts, a likely reflection of the secular liberalization of Chinese sexual mores (Parish et al., 2007). In addition, respondents who grew up in an urban area or resided in a more urban place at the time of the survey, as well as those who resided in the south or east coasts of China at the time of the CHFLS administration, were significantly more likely than others to report having engaged in premarital sexual intercourse.

The final model in Table 2, Model 4, presents the results for whether the respondent tested positive for an STI. Once again, the coefficient for the sex ratio was positive and statistically significant. When faced with a relative abundance of men in their local marriage market at age 20, women were more likely to contract an STI. The apparent effect of the male-to-female sex ratio on the odds that women would test positive for an STI was at least moderate in strength: An addition of 1 man per 100 women translated into a 2.2% increase in the odds that women would test positive for an STI {= [$e^{(.022)}$ − 1] × 100}, and an addition of 12.5 men per 100 women, which corresponds to recent 1985–2005 changes in the sex ratio at birth, translates into an almost one-third increase in the odds of testing positive for an STI (= $\{[e^{(.022)(12.5)}] - 1\} \times 100$).

At a borderline significance level, the likelihood of having an STI was significantly higher for members of the 1960s birth cohort than for members of the 1950s birth cohort and significantly higher for residents of the south and east coastal areas than for residents of other regions. None of the other independent variables was significantly associated with the likelihood of having an STI.

As we suggested earlier, current sexual activity, forced sex, and premarital sex may be risk factors for an STI and may mediate the effect of the sex ratio on this outcome. In Table 3 we examine this possibility by adding these potential mediating factors to Model 4 of Table 2. It was perhaps surprising that none of the coefficients for these variables was statistically significant, and their inclusion reduced the coefficient for the sex ratio by only a modest amount, from .022 (in Model 4, Table 2) to .020 (Table 3). We admit that, absent data on the precise timing of these events and behaviors, we cannot be certain that respondents had recent, premarital, or forced sex prior to contracting an STI, and thus these results must be weighed cautiously. Nevertheless, these findings suggest that other factors, such as extramarital intercourse, the number of different sex partners, or the STI status of male partners may help explain why women were more likely to contract an STI when they were exposed to a relatively abundant supply of men in their local marriage market at age 20.

Additional Analyses

We conducted several sets of supplementary analyses to check on the robustness of our results. One limitation of our analysis is that it was not possible to measure the sex ratio of the community in which the CHFLS respondents lived at age 20 if they migrated into their community after that age. The CHFLS does not record respondents' complete residential histories, which would have allowed us to determine their community of residence—and thus the sex ratio to which migrant women were exposed—at age 20. But the CHFLS did ask respondents whether, at the time of the survey, they were living in the community in which they grew up. As a check on possible problems caused by intercommunity migration, we reestimated all of the models using only respondents who reported residing at the time of the survey in the same community (village, county, city, or prefecture) in which they grew up. This is a stringent check, because these analyses omit female respondents who migrated into their current community before turning age 20 and thus for whom the estimated sex ratio, using our procedures, does represent the appropriate number of men available to them in their young adult years. In models that omit these intercommunity migrants we found the same basic pattern of effects as observed in Table 2 but, in part because of the reduced sample size, the coefficients for the sex ratio were not always statistically significant. Specifically, the coefficient for the sex ratio remained significant for being sexually active ($b = 0.033$, $p < .001$) and for having had premarital sex ($b = 0.016$, $p < .05$), but not for forced sex ($b = 0.012$, $p = .208$) or testing positive for an STI ($b = 0.014$, $p = .293$).

We also estimated models in which we measured the sex ratio by the age-specific numbers of men and women (using 7-year age groups with the 2-year staggering described earlier) at the time of the CHFLS interview. Measuring the sex ratio with this strategy assumes that gender differences in mortality and migration play negligible roles throughout a cohort's life course; that is, this measurement strategy assumes that the number of men to whom women were exposed in the year 2000 (the date of the CHFLS) adequately proxies for the number of men to whom these women were exposed earlier in their lives when they were at risk of experiencing the outcomes. This is probably a reasonable assumption for the younger respondents but may be less tenable for the older respondents. This measurement strategy does have the advantage of drawing entirely on the full-count 2000 census data, thus eliminating the influence of errors in the measurement of the sex ratio incurred by using the 1982 and 1990 1% samples. Moreover, the sex ratio encountered at the time of the CHFLS may be closer in time to the risk period for experiencing some of the events—in particular, current sexual activity and the contraction of an STI—that serve as outcome variables in our analysis. Results from these supplementary analyses were reasonably similar to those we report in Table 2. We observed statistically significant positive effects of the sex ratio on being currently sexually active ($b = 0.049$, $p < .001$), premarital sex ($b = 0.014$, $p < .10$), and having an STI ($b = 0.031$, $p < .05$), but not forced sex ($b = 0.017$, $p = .147$).

Third, we examined whether the associations between the sex ratio and women's sexual outcomes varied across birth cohorts. Members of the older cohorts grew up at a time in which premarital sexual behavior was strongly proscribed (Parish et al., 2007) but in which violence against women was more tolerated (Tang & Lai, 2008; Tang, Wong, & Cheung, 2002). Thus, it is possible that among the older cohorts, premarital sexual activity would be relatively unresponsive to sex ratio imbalances. Given limited freedom to engage in premarital sexual intercourse, older women may have been unlikely to engage in these behaviors even when afforded the opportunity to do so by a relative abundance of men in their communities. Conversely, the likelihood of being sexually victimized may have been comparatively more responsive to sex ratio imbalances among the older than younger cohorts. We explored these possibilities by adding to the regression models shown in Table 2 product terms representing the interaction between the sex ratio and birth cohort. We found no strong evidence that the associations between the sex ratio and women's sexual outcomes vary significantly across birth cohorts. Few of the coefficients for the relevant interaction terms were statistically significant, and their inclusion as a group failed to significantly improve the fit of the models.

Discussion

China has been experiencing a dramatic and growing imbalance between the numbers of women and men in its population, and the increasing surplus of men is thought to influence sexual behavior and the spread of sexually transmitted diseases (Ebenstein & Jennings, 2009; Poston & Glover, 2005; Tucker et al., 2005); nevertheless, few studies have examined this presumed impact of sex ratio imbalances on women's sexual experiences. We addressed this issue by merging individual-level data from the CHFLS with census-derived measures of the numerical availability of men in women's age groups and local communities. We derived hypotheses from demographic-opportunity theory positing effects of the sex ratio on four dimensions of women's sexual outcomes: whether they (a) are sexually active, (b) have experienced forced sex, (c) have had premarital sex, and (d) test positive for an STI. We found statistically significant and moderately strong effects of the sex ratio on all four measures of women's sexual outcomes. Although our findings are not uniformly robust to sensitivity tests, they generally support hypotheses derived from demographic-opportunity theory that posit that the relative supply of men in local communities affects the frequency and form of women's sexual encounters.

Our findings may have implications for social policy insofar as they indicate how China's looming imbalance in the adult sex ratio might alter secular trajectories in women's sexual experiences. We found that a surplus of men was associated with an increase in the odds that women will experience premarital sexual intercourse. China has witnessed substantial ideological and cultural shifts, as well as improvements in women's status, over recent decades. These shifts toward a less traditional social environment are reflected in changes in sexual behaviors, such as more pervasive premarital sexual activity (Parish et al., 2007; Sheng, 2005). Our findings suggest that this secular trend in women's premarital sexual activity may be steeper than what would have occurred in the absence of high sex ratios.

We also found that women are more likely to be forced to have sex when they are numerically scarce relative to men. In China, women's low status has made them particularly vulnerable in both their own households and the larger society (Tang et al., 2002). The growing empowerment of Chinese women may help combat this vulnerability. But at the same time, the positive effect of the sex ratio on women's victimization is problematic given the projected increase in male surplus over coming decades. Any reduction in women's risk of sexual victimization generated by movement toward a less traditional and patriarchal sociocultural climate may be at least partially offset by China's growing surplus of men.

Our results also suggest that China's impending surplus of adult men may lead to an increased risk of HIV and AIDS among women. Rates of HIV, AIDS, and other sexually transmitted diseases have been increasing rapidly in China (Grusky, Liu, & Johnston, 2002; Hong & Li, 2009), but the extent to which a growing abundance of men underlies these trends is not well understood. Our results, indicating a positive association between the sex ratio and the probability that women have contracted an STI, are instructive both because STIs are a risk factor for HIV and AIDS and because of likely similarity in their determinants and modes of transmission (Galvin & Cohen, 2004; Yang et al., 2005).

We acknowledge several limitations to our study. For women who migrated into their communities after age 20, we could not determine definitively the sex ratio of the community in which they lived at age 20. We also were unable to date precisely when in women's lives they experienced premarital sex, forced sex, or contracted an STI. These limitations prevented the estimation of event-history analyses that would treat the sex ratio as a time-varying covariate matched to each period that women were at risk of experiencing the sexual events.

We also note that a key challenge in attempting to infer causal effects of the sex ratio on women's sexual experiences is that the sex ratio may be at least partly determined by forces that might also influence these outcomes. For example, the gender-biased reproductive behaviors that create a surplus of boys and a deficit of girls are also likely to reflect a patriarchal system that devalues the contributions of women, and this devaluation of women may help shape the sexual experiences examined here. Hence, the observed sex ratios may be partly endogenous to the outcomes we examined. Although we controlled for some community-level characteristics that may contribute to a patriarchal cultural system, we cannot be certain that we removed all possible confounding influences.

We also acknowledge that using these findings to project the future of Chinese women's sexual experiences in the face of a growing numerical surfeit of men is a risky undertaking. One important difference between the current cross-sectional intercommunity variation in the adult sex ratio we used and the looming imbalance in adult sex ratios in China's future is that, currently, men can (at least theoretically) move from a community with few women to a community with more women; in the future, however, there will likely be few if any communities with an abundance of women to serve as destinations for potential internal migrants. It is also possible that China may adjust to its projected surplus of adult men through mechanisms such as male emigration or the importation of wives that would temper the impact of these sex ratio imbalances.

Our findings point to several possible directions for future research. That a surplus of men is associated with women's increased likelihood of sexual victimization could mean that, when women are scarce, men exercise greater control and force over women in the society at large, that they maintain stricter control of their current partners in the private sphere, or both. Our measure of forced sex does not distinguish among physical coercion used by husbands, partners, or strangers; neither did we have information on the dynamics of interpersonal relationships. Future research should explore the underlying dyadic processes involved in how sex ratios alter men's and women's relative power.

Future research might also profit by exploring how the growing surplus of men (and the attendant deficit of women) will affect the broader nature of social relations between men and women, and women's status more generally, in China. On the one hand, women may attain greater power in society when they are in short supply because men will need to offer more resources in order to strike a marital bargain (Guttentag & Secord, 1983). On the other hand, under conditions of female scarcity, men may also more vigorously guard their mates and limit their participation in the broader society (South & Trent, 1988). How these disparate processes balance out may be important for understanding more clearly the relationships between population sex ratios and family-related behavior. In China, as well as elsewhere in the world, gender inequality is likely tied to mate availability in ways that have yet to be thoroughly investigated.

Acknowledgments

We thank Yong Cai, Baochang Gu, Lin Guo, Zai Liang, Kelly McGeever, Jeremy Pais, William Parish, Carey Sojka, and Michelle Zagura for assistance. This research was supported by a grant from the National Institute of Child Health and Human Development (R21 HD057289). The Center for Social and Demographic Analysis of the University at Albany also provided technical and administrative support for this research through a grant from the National Institute of Child Health and Human Development (R24 HD044943).

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Table 1

Weighted descriptive statistics for variables used in analysis of sexual experiences: Women ages 20–44, Chinese Health and Family Life Survey (*N* = 1,369)

Variables Dependent	**Description**	**%**	***n***
Sexually active	R reported engaging in sexual intercourse in past year	89.04	1,369
Forced sex	R reported having been forced to have sex against her will	6.65	1,338
Premarital sex	R reported having had sexual intercourse prior to marriage	12.38	1,369
Has STI	R’s urine test was positive for gonorrhea, chlamydia, or trichomonas	4.49	1,241
Independent			*SD*
Sex ratio	Estimated number of men ages 17–23 per 100 women ages 15–21 in R’s community when R was age 20	108.51	12.25
Education			
Never attended school	R never attended school	0.11	
Elementary school	R’s highest schooling is elementary school	0.29	
Junior high school	R’s highest schooling is junior high school	0.40	
Senior high school	R’s highest schooling is senior high school	0.15	
Junior college	R’s highest schooling is junior college	0.03	
University/graduate school	R’s highest schooling is university or graduate school	0.01	
Birth cohort			
1950	R was born 1950–1959 (1 = yes)	0.14	
1960	R was born 1960–1969 (1 = yes)	0.44	
1970	R was born 1970–1979 (1 = yes)	0.42	
Urban residence at age 14	R resided in urban area at age 14 (1 = yes)	0.15	
South/east coast	R’s community is on the south or east coast (1 = yes)	0.11	
Community education	Proportion of community population with a senior high school education	0.29	0.15
Community urbanization	Proportion of community population residing in urban village or neighborhood	0.45	0.37

Note. Descriptive statistics for independent variables are taken from the largest sample (*N* = 1,369). Standard deviations for dummy variables are not shown. R = respondent; STI = sexually transmitted infection.

Table 2

Multilevel logistic regression analyses of sexual experiences: Women ages 20–44, Chinese Health and Family Life Survey (N=1,369)

	Model 1 Sexually active			**Model 2 Forced sex**			**Model 3 Premarital sex**			**Model 4 Has STI**		
	b	***SE***	e^x	***b***	***SE***	e^x	***b***	***SE***	e^x	***b***	***SE***	e^x
Sex ratio	0.038**	0.008	1.039	0.018*	0.009	1.018	0.017*	0.007	1.017	0.022*	0.011	1.022
Education												
Never attended school (ref.)												
Elementary school	−0.179	0.813	0.836	0.945	1.081	2.573	−0.944†	0.522	0.389	0.953	1.066	2.593
Junior high school	−0.771	0.774	0.463	2.060*	1.036	7.846	0.007	0.475	1.007	0.965	1.050	2.625
Senior high school	−0.708	0.786	0.493	1.887†	1.048	6.600	−0.393	0.492	0.675	0.886	1.065	2.425
Junior college	−1.082	0.813	0.339	2.537*	1.069	12.642	−0.522	0.537	0.593	−0.654	1.280	0.520
University/graduate school	−0.744	0.888	0.475	2.650*	1.131	14.154	0.097	0.599	1.102	−0.130	1.471	0.878
Birth cohort												
1950 (ref.)												
1960	0.637†	0.334	1.891	−0.285	0.252	0.752	0.404†	0.232	1.498	0.731†	0.385	2.077
1970	−1.979**	0.288	0.138	−0.824**	0.283	0.439	0.807**	0.232	2.241	0.428	0.402	1.534
Urban residence at age 14	−0.419†	0.214	0.658	−0.258	0.232	0.773	0.367*	0.174	1.443	0.389	0.290	1.476
South/east coast	0.069	0.192	1.071	−0.822**	0.228	0.440	0.883**	0.160	2.418	0.803**	0.266	2.232
Community education	−1.614	1.132	0.199	−0.656	1.273	0.519	−1.471†	0.838	0.230	0.490	1.467	1.632
Community urbanization	0.318	0.552	1.374	1.086†	0.609	2.962	0.948*	0.424	2.581	−0.893	0.667	0.409
Constant	0.008	1.122		−5.840**	1.427		−4.336**	0.876		−6.468**	1.597	
n persons		1,369			1,338			1,369			1,241	
n communities		37			37			37			37	
Log likelihood		−461.046			−404.610			−625.365			−290.997	

Note. STI = sexually transmitted infection.

† $p < .10$.

* $p < .05$.

** $p < .01$. (two-tailed tests)

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Table 3

Multilevel logistic regression analysis of having a sexually transmitted infection (STI): Women Ages 20–44, Chinese Health and Family Life Survey (n =1,221)

	Has STI		
	b	***SE***	***e^x***
Sex ratio	0.020†	0.012	1.020
Sexually active	0.179	0.418	1.196
Forced sex	0.496	0.352	1.642
Premarital sex	−0.063	0.292	0.939
Education			
Never attended school (ref.)			
Elementary school	0.932	1.067	2.540
Junior high school	0.925	1.051	2.522
Senior high school	0.852	1.066	2.344
Junior college	−0.688	1.282	0.503
University/graduate school	−0.243	1.478	0.784
Birth cohort			
1950 (ref.)			
1960	0.737†	0.386	2.090
1970	0.517	0.416	1.677
Urban residence at age 14	0.409	0.291	1.505
South/east coast	0.827**	0.277	2.286
Community education	0.511	1.474	1.667
Community urbanization	−0.917	0.675	0.400
Constant	−6.489**	1.664	
n persons	1,221		
n communities	37		
Log likelihood	−289.044		

$^{\dagger}p < .10$.

$^{**}p < .01$. (two-tailed tests)