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Sexually-Integrated Workplaces and Divorce: Another Form of On-the-Job Search

Terra McKinnish
Department of Economics, University of Colorado at Boulder
Boulder, Colorado

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Center for Economic Analysis
Department of Economics

University of Colorado at Boulder
Boulder, Colorado 80309

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by

Terra McKinnish
Department of Economics, UCB 256
University of Colorado-Boulder
Boulder, CO 80309-0256
mckinnis@Colorado.edu
303-492-6770

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Abstract

As women have entered the work force and occupational sex-segregation has declined, workers experience increased contact with the opposite sex on the job. Because this contact lowers the cost of search for alternative mates, the sex-mix a worker encounters on the job should affect the probability of divorce. This paper uses 1990 Census data to calculate the fraction of workers that are female by industry, occupation and industry-occupation cell. These results are then used to predict divorce among ever-married respondents in the 1990 Census. Two separate strategies are employed to address endogenous occupation and industry choice. In the first, sex-mix measures are calculated at the state level so that industry and occupation fixed-effects can be included in the regression. In the second, the sex-mix a worker faces on the job is instrumented with the industry and occupation composition of employment in the worker's local area. The results indicate that those who work with a larger fraction of workers of the opposite sex are more likely to be divorced.

1. Introduction

In discussing the economics of marriage and divorce, Becker (1991) points out that imperfect information at the time of marriage and the acquisition of additional information while married is a key determinant of divorce. He states:

> "Imperfect information can often be disregarded without much loss in understanding, but it is often the essence of divorce . . . participants in marriage markets hardly know their own interests and capabilities, let alone the dependability, sexual compatibility and other traits of potential spouses. Although they date and search in other ways to improve their information, they frequently marry with highly erroneous assessments, then revise these assessments as information improves after marriage."(p.324)

Information acquired during marriage can change both an individual's assessment of the quality of their current spouse as well as their assessment of their "outside alternatives."

As the labor force participation of women has increased and as women have increasingly found employment in industries and occupations that were once almost exclusively male, on the job contact with members of the opposite sex has increased. This substantial increase in workplace interaction between men and women is a major change in our society that has largely been ignored by economists. One important consequence of this workplace contact is that it allows married men and women to acquire additional information about their outside alternatives at a much lower cost.

This paper examines the extent to which the sex-mix an individual encounters on the job affects their marital status. Specifically, the 1990 Census is used to calculate the fraction of workers that are female for each industry, occupation and industry-occupation cell. These results are then used to predict divorce among ever-married respondents in the 1990 Census. Choice of occupation and industry is endogenous and could be related to other unobserved characteristics of individuals that make them more or less prone to divorce. Two separate strategies are employed to address this unobserved heterogeneity. In the first sex-mix measures are calculated at the state level, allowing the inclusion of occupation and industry fixed-effects in the regression model. In the second, the sex ratio a worker faces on the job is instrumented with the industrial and occupational composition of employment in the worker's local labor market. The results indicate that women who work with a larger fraction of male coworkers are more likely to be divorced, and, to a lesser extent, men who work with a larger fraction of female coworkers are more likely to be divorced.

It has long been argued that the increased labor force participation of women was a major factor in the rise in divorce rates during the second half of the 20th Century. As Cherlin (1992) writes, "As for the rise in divorce and separation, almost every well-known scholar who has addressed this topic in the twentieth century has cited the importance of the increase in employment of women."[1] The usual causal mechanism cited for this relationship is that the increase in labor market opportunities increased women's income, and therefore utility, outside of marriage. It is less recognized, however, that part of the effect of female employment on divorce operates through the increased interaction of men and women in the workplace.

2. Literature Review

In light of the dramatic rise in the divorce rate since World War II, there is a large literature that attempts to explain the increased prevalence of divorce. Much of this literature indicates that the rising labor market opportunities of women are at least partially responsible (see Ross and Sawhill, 1975; Michael, 1988; Greenstein, 1990; McLanahan, 1991; Cherlin, 1992; Ruggles, 1997; and South, 2001). For the most part, these studies do not consider the effect of female employment on workplace contact between men and women. One exception is South (2001) who studies changes in the effect of wives' employment on divorce over time using

[1] Also cited in Ruggles (1997).

the Panel Study of Income Dynamics (PSID). South cites the fact that "historical declines in occupation sex segregation . . . have likely meant that more and more employed married women work in close proximity with men who might serve as more attractive mates than their current husband," as a reason that the effect of wives' employment on divorce might have changed over time, although he does not directly test this hypothesis.

The literature that relates the marital status of individuals to the proximity of potential mates is largely limited to the study of the relationship between the probability an individual marries and the supply of potential spouses in the state or local geographic area. Much of this literature focuses on racial differences in marriage rates and is motivated by the contention of Wilson (1987) that marriage rates for black women are low relative to white women because of the limited supply of employed black men available as potential spouses (e.g. Lerman, 1989; Olsen and Farkas, 1990; Fitzgerald, 1991; Lichter, LeClere and McLaughlin, 1991; Brien, 1998).

The question of whether the availability of alternative spouses affects divorce rates has received considerably less attention. South and Lloyd (1995) consider whether the supply of alternative spouses in the local geographic area affect the probability of divorce. They find that divorce is more common in areas where the ratio of unmarried men to unmarried women is either very high or very low. They also find that divorce is more likely in areas with a high female employment rate, even controlling for the employment and earnings of the wife. This latter result clearly has numerous interpretations, but the authors posit that increased labor force participation of women increases social interaction of men and women in the workplace.[2] While both South and Lloyd (1995) and South (2001) theorize that workplace interaction between men

[2] For example, this empirical relationship could simply indicate that an area with a stronger labor market for women provides them with the economic independence to divorce. Alternatively, both the higher rates of labor force participation and the higher rates of divorce might indicate that the local area is more socially progressive.

and women destabilizes marriage, to date that has been no direct test that relates marital dissolution to the actual sex-mix men and women encounter on the job.

In contrast, the theoretical literature provides ample reason to suspect that sex-integration in the workplace will increase the prevalence of divorce. Becker, Landes and Michael (1977), Mortensen (1988) and Chiappori and Weiss (2001) all apply search theory to marriage and divorce decisions, often comparing them to the more familiar job search and on-the-job search for alternative employment. Within this framework, it quickly becomes clear that to the extent that sexual integration in the workplace lowers search costs so that married individuals may more easily meet alternatives mates, divorce rates should increase.[3]

Becker, Landes and Michael (1977) state, "When remarriage is possible, continued marital search may be quite rational," but note that, "marital status often severely limits the effort they can devote to search" (p.1155). Mortensen (1988) also assumes that search is more costly when married than when single. This difference in search costs between the married and single state is why sex integration is particularly salient to the divorce, as opposed to the marriage, decision. It is true that sexual integration in the workplace should also affect the ability of never-married individuals to find mates, but the important realization is that for individuals who are married, search for alternative mates outside of the workplace is extremely limited and very costly compared to that for single individuals.[4] Therefore, while sex integration in the workplace almost certainly effects the marriage decision, the larger effect should be on the divorce decision.[5]

[3] Fair (1978) develops a model of extramarital affairs, but his is a standard household consumption model, in which non-market time is divided into time spent with family and time spent apart from family in an elicit relationship. As such, search costs do not figure into his model. It is assumed that the alternative partner is readily available.

[4] Lauman et al (1994) report that about 15 percent of individuals met their spouse through work.

[5] Alternatively, if sexual integration in the workplace lowers search costs for single individuals as much as it does for married individuals, the effect on divorce would be ambiguous. The lowered search costs for singles should result in higher-quality matches that are less prone to divorce. This is then an empirical question.

It is important to point out that workplace contact with members of the opposite sex can result in divorce through multiple mechanisms. The first and most obvious is that an individual finds a potential spouse at work that is more appealing than their current mate, and divorces in order to marry that person. The second is that workplace contact leads to an extra-marital affair that disrupts the marriage even if the liaison was unlikely to result in a long-term relationship or marriage. The final mechanism is less obvious, because it does not require the development of an actual romantic relationship with a coworker. The mere fact that an individual is meeting a lot members of the opposite sex at work may change their perceptions of their outside alternatives, causing them to feel less satisfied with their current partner and more likely to divorce.[6] Both Udry (1981) and White and Booth (1991) find evidence in survey data that individual's perceptions of their ability to replace or improve upon their mate is a significant predictor of divorce, even controlling for measures of marital satisfaction.

One important additional insight from the theoretical literature on divorce is the finding of Chiappori and Weiss (2001) that there is plausibly a feedback mechanism that causes marriage market to be highly sensitive to exogenous shocks. The basic idea is that "random search process creates a *meeting externality* where by one divorce (marginally) increases the remarriage probability of other divorcees" (p.20). If the increased sexual integration of the workplace increases the number of divorces, this in turn increases the rate at which individuals come into contact at work with members of the opposite sex who are not married. Assuming search is more fruitful when there is a larger supply of single individuals, this in turn should increase the returns to search for those who are still married and further increase the divorce rate. This suggests that a relatively small initial decline in search costs for married individuals could have a rather large effect on the divorce rate.

[6] A similar point is made by South and Lloyd (1995).

3. Empirical Analysis

A. Sex-Segregation by Occupation and Industry

The economics literature has already documented that men and women in the workplace are heavily segregated by occupation and, to a lesser extent, by industry. This feature of the labor market has generally been of most interest to those researchers attempting to explain the gap between male and female wages (e.g. Bayard et al, 2000; Macpherson and Hirsch, 1995; and Sorenson, 1990). This literature also documents the declines in occupational segregation over time. For example, using CPS data, Macpherson and Hirsch report that in 1973 the average female worker worked in an occupation that was 72.1 percent female and the average male worker worked in an occupation that was 17.6 percent female. In 1993, the corresponding statistics were 68.2 and 28.8 percent.

For this analysis, sex-segregation measures are calculated using the 1990 Census for each of 235 civilian industries, 501 civilian occupations and 52,709 civilian industry-occupation cells. The statistic of interest is the fraction of workers between the ages of 18 and 55 that are female. About half of the industry-occupation cells have no more than 5 observations. These industry-occupation combinations are omitted from the analysis because so few observations are available to calculate the fraction of workers that are female, leaving 26,694 industry-occupation cells.

Distributions of these statistics for men and women are reported in Table 1. It is clear that there is still substantial segregation by industry and occupation. The median woman works in an occupation that is 74% female and an industry that is 61% female, while the median man works in an occupation that is 26% female and an industry that is 32% female. The distributional statistics, however, indicate that there is substantial variation in the sex-mix experienced by men and women on the job. For example, about a quarter of women work in occupations that are at least 50 percent male, while a quarter of men work in occupations that are at least 40 percent female.

Table 2 presents some preliminary evidence on sex segregation and divorce. The table categorizes men and women based on whether the percent female in their industry-occupation cell is less than 25 percent, between 25 and 49 percent, between 50 and 74 percent, or 75 percent or more. Among the women, there is a very visible relationship between percent female and divorce. Only 5.8 percent of women work in industry-occupation combinations that are less than 25% female, but their divorce rate is 24.2 percent. In contrast, 56.2% of women work in industry-occupation combinations that are at least 75 percent female, but their divorce rate is only 17.8 percent. For men, there is a slight positive relationship between percent female in industry-occupation, but it is less pronounced.

The 1990 Census data are also used to calculate the fraction of non-institutionalized residents of each Public Use Microdata Area (PUMA) between the ages of 18 and 55 that are female and the fraction of men and women ages 18 to 55 in each PUMA that are employed. South and Lloyd (1995) find a curvilinear relationship between local sex-ratios and divorce, so that divorce is more likely when either there is an oversupply of women or an oversupply of men. A similar relationship is modeled in the regression analysis below using linear and squared terms.

B. Sample of Analysis

The sample from the 1990 Census used in the regression analysis includes all non-institutionalized, ever-married, non-widowed individuals ages of 18 and 55 who report an industry and occupation.[7] Individuals are dropped from the sample if they report an industry-occupation cell for which no more than 5 observations are available for calculating the fraction

[7] Respondents for whom marital status, industry or occupation are allocated are omitted from the sample.

of workers who are female.[8] The final sample consists of 1,937,790 women and 1,925,662 men. Descriptive statistics are reported in Table 3.

One concern about the sample is that only those individuals who have worked within the past 5 years will report an industry or an occupation in the Census data. In the sample of non-institutionalized ever-married women ages 18-55, 14.8 percent of married women do not report an industry or occupation and 9.2 percent of divorced women similarly must be excluded from the sample. For the sample of men, 1.8 percent of married men and 5.1 percent of divorced men do not report an occupation or industry. The sample used in the analysis conditions on a certain level of labor force attachment, which can be endogenously determined by marital status.

C. OLS Analysis

The baseline regression model used is the linear probability model:

$$\begin{aligned} Y_{ionls} &= \beta_0 + \beta_1 FrFemOCC_o + \beta_2 FrFemIND_n + \beta_3 FrFemLOC_l + \beta_4 (FrFemLOC_l)^2 \\ &\quad + \beta_5 FrMEMP_l + \beta_6 FrWEMP_l + X_{ionls}\beta_7 + STATE_s \delta + (STATE_s * Urban_i)\phi + \varepsilon_{ionls} \end{aligned} \tag{1}$$

Where for person i working in occupation o and industry n, living in local PUMA l and state s, Y is an indicator for divorce, *FrFemOCC* is the fraction of workers in the occupation that are female, *FrFemIND* is the fraction of workers in the industry that are female, *FrFemLOC* is the fraction of residents ages 18-55 of the PUMA that are female, *FrMEMP* is the fraction of men employed in the PUMA, *FrWEMP* is the fraction of women employed in the PUMA, X is a vector of individual control variables, *STATE* is a vector of state indicator variables and *STATE*Urban* is a state-urban fixed effect. These two sets of fixed-effects control for unobserved differences across states and differences between urban and rural areas within states. The individual controls include age, age-squared, race indicators (black, asian, other), a Hispanic

[8] This omits 39,991 individuals, a little less than 1 percent of the sample.

ethnicity indicator, an urban residence indicator, and education indicators (high school degree, some college, college degree, more than college degree).

The cross-sectional nature of the data raises the concern that because we only observe marital status at one point in time, we only observe those individuals who are currently divorced. We have no way of knowing if an individual has divorced and remarried. But to the extent that workplace contact, through the mechanisms discussed above, generates divorce that is not immediately followed by remarriage, the effect of interest can be identified in the cross-sectional census data.[9]

The initial regression results are reported in Table 4. Columns 1-3 report the results for women and columns 4-6 report the results for men. The first and fourth columns report the results obtained from estimating the regression model specified in equation (1). For women, working in industries and occupations with a higher fraction female lowers the probability of divorce. For men, working in an occupation with a higher fraction female raises the probability of divorce, but the fraction female in the industry of employment has no effect. In columns 2 and 5, the two variables for fraction female in occupation and industry are replaced with the fraction female in the industry-occupation cell. The results show that that women working in an industry-occupation cell with a higher fraction of women are less likely to be divorced, while men working in an industry-occupation cell with a higher fraction of women are more likely to be divorced.

[9] Kreider and Fields (2001) report that in 1996, 40% of men who had ever divorced were still divorced and 45.5% of women who had ever divorced were still divorced. For those that had remarried, the median time to remarriage was 3.3 years for men and 3.1 years for women. Norton and Miller (1992) report results from a 1990 survey indicating that the vast majority of those who divorce eventually remarry. Of those women in their survey who had divorced and remarried, the median duration to remarriage was 2.5 years (the 25th percentile was 1 year and the 75th percentile was 5 years).

The magnitude of the effect for men is modest, but the effect for women is fairly substantial. For example, a woman moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .540 to .928 would decrease her probability of divorce by 3.2 percentage points. A man moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .072 to .404, would increase his probability of divorce by 0.78 percentage points.

There are any number of reasons that the effect of sex-mix in occupation and industry could differ between men and women. For example, if the amount and nature of contact with coworkers differs between jobs in male-dominated occupations and jobs in female-dominated occupations, a job that is 75% female may not have the same effect on search costs for men as a job that is 75% male has on search costs for women. This difference in search costs could also reflect differences in behavior between men and women in approaching members of the opposite sex at work.

In columns 3 and 6, variables measuring the fraction female in the occupation, industry and industry-occupation cell are all included in the regression at the same time. For both men and women, the effect of the fraction female in the industry-occupation cell remains with the predicted sign and the effect increases in magnitude. The independent effects of fraction female in the industry and the occupation, however, reverse signs in three of the four cases. These effects are difficult to interpret. They suggest, for example, that a man working in an industry-occupation cell with a high percentage of female co-workers is more likely to get a divorce, but he is less likely to get a divorce if there are more women in his industry and occupation outside of his actual industry-occupation combination.[10]

[10] One issue is that the correlation between the industry-occupation measure and the occupation measure is .90, which makes it somewhat difficult to identify independent effects. For women, regressing *FrFemOCC* on the other

The other coefficients reported in Table 2 are for the PUMA-specific variables. As expected, there is a U-shaped relationship between the probability of divorce and the percent of women in the local PUMA. Also as one would expect, a higher employment rate for men in the local area is associated with a lower probability of divorce and a higher employment rate for women in the local area is associated with a higher probability of divorce.

One concern about the results in Table 4 is that higher divorce rate for women working in occupations and industries with more men may reflect the fact that wages tend to be higher in these types of jobs, in which case a simple bargaining model would predict that women in these occupations and industries would have higher rates of divorce. Therefore, in Table 5, wage controls by industry, occupation and location are be added to the regression analysis.

For each occupation, industry, industry-occupation cell and PUMA, the mean and variance are calculated separately for male and female wages.[11] The mean male and female wages and the logarithms of male and female wage variances are then included as controls in the divorce regressions. Descriptive statistics for the wage measures are reported in Appendix Table A1.[12] The regression results are reported in Table 5. The patterns in the coefficient estimates for percent female in occupation, industry and industry-occupation are very similar to those reported in Table 4, but the magnitudes have changed. For example, the magnitude of the effect of fraction female in industry-occupation cell reported in columns 2 and 5 has increased about 14 percent for women, but has decreased over 60 percent for men.

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right-hand side variables used in column 3 generates an R-squared statistic of .98. For men the corresponding R-squared statistic is .93.

[11] Workers with wages less than $2/hr or greater than $200/hr are excluded from the wage calculation.

[12] The samples are somewhat reduced, as regressions do not include those workers in industry-occupation cells that do not contain at least two male workers and two female workers with wages between $2 and $200/hr do not have sufficient information to calculate wage dispersion measures by gender. The sample means for these reduced samples, however, are almost identical to those reported in Table 3.

Because the addition of these wage controls has little effect on the coefficient estimates for the PUMA-specific variables, the coefficient estimates for these location controls are omitted from the table for brevity. The coefficient estimates for the occupation, industry and industry-occupation wage controls are reported in Table 5. It is difficult to predict the effects of these wage measures on divorce, because multiple mechanisms are at work. For example, if a man works in an occupation or industry with above-average wages, this suggests that his earnings potential is also above average.[13] This would tend to make his marriage more stable to the extent that his current spouse should value their marriage more highly. On the other hand, the higher wage also makes him more attractive to potential alternative spouses. These results are therefore not a primary focus of the paper.

The general finding in Table 5 is that, higher wages lower the probability of divorce, with the exception of higher occupational wages for men, which lower the odds of female divorce. Interestingly, higher wage dispersion tends to have a negative effect on divorce for women and a positive effect on divorce for men, although there exceptions in both cases.

D. Fixed-Effects Analysis

Choice of occupation and industry is potentially endogenous. One might argue, for example, that women that enter male-dominated occupations are more independent and less family-oriented and will be more prone to divorce regardless of exposure to alternative mates. It should be noted that the selection could work in the opposite direction. Women who work in male-dominated occupations tend to have higher educational attainment. Education is negatively correlated with divorce, probably due in part to the fact that these women delay marriage to later

[13] Individual wages are not included as controls, as these could obviously be endogenous to marital status. The same goes for fertility related measures.

ages, which should increase the quality of the match.[14] Therefore, it is also possible that women who work in more sexually-integrated occupations have unobserved characteristics that make them less, not more, prone to divorce.

One potential solution is to include industry and occupation-specific fixed-effects in the model. This can be accomplished by calculating the sex-mix measures for industry, occupation and industry-occupation cell at the state, rather than national level. This provides the necessary additional variation in the sex-mix measures so that the sex-mix effects are identified when industry and occupation fixed-effects are added as controls.[15]

The main weakness of this approach is that the cross-state variation in sex-mix within occupation or industry category will be substantially less than the cross-occupation or cross-industry variation in sex-mix. Some diagnostic regressions can shed some light on the amount of variation in the sex-mix measures available to identify the effect of interest. A regression of state-level sex-mix in occupation on occupation fixed-effects generates an R^2 statistic of 0.988 for men and 0.997 for women. A regression of state-level sex-mix in industry on industry fixed-effects generates an R^2 statistic of 0.993 for men and 0.996 for women. A regression of state-level sex-mix in industry-occupation cell on occupation and industry fixed-effects generates an R^2 statistic of 0.922 for men and 0.981 for women. These results suggest that to the extent this

[14] For the regression models estimated in Table 5, the unreported coefficient estimates for the education variables indicate that, conditional on marriage, women with a high school diploma but less than a college degree have a 1.7-3.5 percentage point lower probability of divorce than those without a high school diploma and those with at least a college degree have a 5.6-8.9 percentage point lower probability of divorce than those without a high school diploma. For men, the correspondent results are .3 to 2.2 percentage points lower probability of divorce for those with a high school diploma but no college degree and 4.7-5.3 percentage points lower probability of divorce for those with at least a college degree.

[15] Because of the requirement that there be more than five workers used to calculate sex-mix and wage statistics for industry-occupation cell, calculating these measures at the state level reduces the sample by 28% for men and by 24% for women. The results are not particularly sensitive to this reduction in sample or to the switch to state-specific sex-mix measures. Re-estimating the models in columns 2 and 5 of Table 5 on this reduced sample using the state-specific measures reduces the coefficient on sex-mix in industry-occupation cell from –0.0942 to –0.0806 for women and increases it from 0.0091 to -.0110 for men. Therefore, most of the difference in coefficients between Tables 5 and 6 is due to the inclusion of occupation and industry fixed-effects.

approach is viable, it is most viable for the results using sex-mix in industry-occupation category, and it is more viable for men than for women.

The results of the fixed-effects analysis are reported in Table 6.[16] The results for women are again reported in columns 1-3. Looking across all three sets of results, the coefficient on fraction female in industry-occupation cell has the anticipated negative sign. The magnitude of the coefficients is smaller than what is estimated in Table 5. In fact, as discussed further below, including the fixed-effects brings the magnitudes of the effects for men and women much closer together. The coefficients on fraction female in occupation and fraction female in industry, however, are either positive or insignificant. The results for men are again reported in columns 4-6. All coefficients on the sex-mix measures have the anticipated positive signs, and the magnitudes are larger in magnitude than those reported in Table 5. The sex-mix measures for which there is the most available variation independent of the fixed effects, those for men and those for industry-occupation cell, are the ones that have the coefficient estimates that are most consistent with expectations.

The magnitude of the coefficient on fraction female in industry-occupation cell reported in column 2 is such that moving a woman from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell would decrease her probability of divorce by 1.4 percentage points, or 7.2% of the mean female divorce probability of 0.194. Moving a man from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell would increase his probability of divorce by 1.0 percentage points, or 7.5% of the mean male divorce probability of 0.133. These results suggest relatively similar effects of sex-mix on divorce for men and women after controlling for unobserved occupation and industry characteristics.

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[16] Due to computer memory constraints, the results in Table 6 are then estimated using a 90% sample of the eligible sample described in footnote 15.

For the fixed-effects analysis in Table 6, wage and wage dispersion characteristics for industries and occupations are calculated at the state level as well. Therefore, it is possible to report coefficient estimates for these variables. The results for women differ from the OLS results in that higher mean wages in occupation, industry or industry-occupation cell are now associated with a higher, rather than lower, probability of divorce. Higher wage dispersion is still associated with a lower probability of divorce for women. The wage results for men are more mixed, with both positive and negative effects estimated for the wage and wage dispersion measures.

E. IV Analysis

As another approach to addressing the endogeneity of occupation and industry choice, the sex-ratio a worker faces in his or her occupation or industry is instrumented with the industrial and occupational composition of employment in the worker's local labor market. For a male worker in PUMA *l*, the instrument for the fraction employment in a worker's occupation that is female is:

$$IVOCCMale_l = \sum_o \frac{ShareMaleEmp_{ol} * FrFemOCC_o}{TotalMaleEmp_l}, \tag{2}$$

where $ShareMaleEmp_{ol}$ is the fraction of total male employment in PUMA *l* that occurs in occupation *o*, $FrFemOCC_o$ is the fraction of *national* employment in occupation *o* that is female, and $TotalMaleEmp_l = \sum_o ShareMaleEmp_{ol}$ is total male employment in PUMA *l*. An analogous instrument can be calculated for the fraction female in a male worker's industry:

$$IVINDMale_l = \sum_n \frac{ShareMaleEmp_{nl} * FrFemIND_n}{TotalMaleEmp_l}, \tag{3}$$

where $ShareMaleEmp_{nl}$ is the fraction of total male employment in PUMA l that occurs in industry n and $FrFemInd_n$ is the fraction of national employment in industry n that is female,

The instruments for a female worker in PUMA l are:

$$IVOCCFem_l = \sum_o \frac{ShareFemEmp_{ol} * FrFemOCC_o}{TotalFemEmp_l}, \tag{4}$$

and:

$$IVINDFem_l = \sum_n \frac{ShareFemEmp_{nl} * FrFemIND_n}{TotalFemEmp_l}. \tag{5}$$

These instruments are calculated for each of the 1725 PUMAs in the 1990 PUMS. Analogous instruments were calculated using share of male and female employment in industry-occupation cell. As will be discussed below, these industry-occupation instruments prove to be less useful in analysis. Industry and occupation fixed-effects are not included in the instrumental variables specification. The instruments do not have sufficient power in the first stage when these fixed-effects controls are added.

These instruments have the appeal that they should be substantially less correlated with individual characteristics than individual's own choice of occupation and industry. Additionally, if workers who are already divorced or generally less committed to marriage seek out employment in more sexually-integrated workplaces, they might respond endogenously to cross-state differences in sex-mix by industry and occupation. Industry and occupation fixed-effects do not address this form of endogeneity. It is the case, however, that areas that have large shares of employment in industries and occupations that tend to be more integrated may differ in social attitudes from places with large shares of employment in industries and occupations that tend to be highly segregated by sex. To the extent that that are unobserved PUMA-specific confounders, the instrumental variables results can still suffer from bias due to unobserved heterogeneity. It should be noted, however, that the regressions do control for state fixed-effects and state-urban fixed-effects. Therefore, the effect of interest is not identified from comparing divorce in Idaho to divorce in California. Nor is it identified from comparing rural central Pennsylvania to Philadelphia. The relevant variation in the instruments is within state and within urban/rural classification. Unobserved heterogeneity in PUMA characteristics should therefore be less problematic.

Instrumental variables results are reported in Table 7. Column 1 instruments the fraction female in occupation and industry encountered by female workers with the variables described in equations (4) and (5). Column 3 instruments the fraction female in occupation and industry encountered by male workers with the variables described in equations (2) and (3).

For the women, the fraction female in their occupation, as predicted by the occupational composition of the workforce in their PUMA, has the predicted negative effect on divorce. The magnitude of the effect is substantially larger than that reported in Tables 4-6. The predicted fraction female in industry, however, has an unexpected positive effect on divorce, although the magnitude is smaller than that for fraction female in occupation. For men, there is a similar pattern in the results. The predicted fraction female in their occupation has the expected positive effect on divorce and is much larger in magnitude than the effects reported in Tables 4-6. The predicted fraction female in the industry, however, has a negative, although smaller, effect.

In columns 2 and 4, the fraction female in the worker's industry-occupation cell is instrumented with both the occupational and industrial composition variables.[17] In both cases the

[17]Even when predicting the fraction of workers in the industry-occupation cell that are female, it was found that the instruments described in equations (2)-(5) performed better than instruments based on the share of employment in industry-occupation cell. For women, the first-stage partial F-statistic on IVOCCFem and IVINDFem is 3220 where the first-stage partial F-statistic on IVIOFem is 621. For men, the analogous F-statistics are 2145 and 942. For men, the weaker instrument produces a moderately smaller coefficient estimate (.0987 vs .1538). For women, the coefficient estimate obtained with the weaker instrument is substantially smaller in magnitude (-.0861 vs -.3247).

results are of the predicted sign and the effects are larger in magnitude than those obtained with OLS estimation. To give an idea of the size of these effects, a woman moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .540 to .928 would decrease her probability of divorce by 12.6 percentage points. A man moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .053 to .404, would increase his probability of divorce by 5.7 percentage points. These are very sizeable effects. Unlike the fixed-effects results, but in keeping with the OLS results, these estimates imply a larger effect of sex-mix on divorce for women than for men.

While in theory it is possible to estimate instrumental variables regressions that include all three of the sex-ratio variables, in practice the instrument based on employment shares in industry-occupation cells does not have sufficient variation independent of the instruments described in equations (2)-(5). It is, therefore, not feasible to estimate instrumental variables versions of the results that appear in columns 3 and 6 of Table 5.[18]

E. Age and Race Specific Results

We might expect the effect of sex-mix on divorce to vary by age and race. Table 8 reports OLS and IV coefficient estimates for the fraction female in industry-occupation cell for age and race-specific sub-samples. The first row merely repeats the results for the full sample that are reported in columns 2 and 5 of Table 5 and columns 2 and 4 of Table 7. The second row reports the results for women and men ages 18-29. The sex-mix measures, the wage and wage dispersion measures, and the location-specific instruments have all been re-calculated on this sample of young adults so that all of the control variables have been recalculated to be specific to

[18] For both men and women, a regression of the occupational structure instrument described in equation (2) on the industrial structure instrument described in equation (3), the analogous industrial-occupational structure instrument and the additional controls used in columns 3 and 6 of Table 5 produces an R-square of .999. As a result, estimating the instrumental variables versions of columns 3 and 6 in Table 5 produce nonsensical results with very large coefficients, extremely large standard errors and p-values greater than .9.

this age group. The coefficient estimates therefore indicate the effect of the fraction of 18-29 year old workers in industry-occupation cell that are female on the probability of divorce among individuals ages 18 to 29. Independent variables and instruments were similarly recalculated for the other age and race specific sub-samples in the table. Overall, the results indicate that the effects are larger for workers 30 and over than young adults, the one exception being the IV results for men aged 30-40. Not surprisingly, the effects are substantially stronger for whites than non-whites. The OLS results for non-whites were re-estimated adding controls for and interactions with the fraction of workers that are non-white, but still the effects of sex-mix on divorce were very small.

5. Conclusions

This paper presents evidence that the fraction of workers in an individual's occupation or industry-occupation combination that are female affects the probability an individual is divorced. Women who work with more men are more likely to be divorced and men who work with more women are more likely to be divorced. The results are more consistent for industry-occupation cell and for occupation than for industry. It could be that the fraction of female workers in an individual's occupation is a better indicator of the amount of workplace contact with members of the opposite sex than the fraction of female workers in an individual's industry.

The results for women indicate that moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell decrease the probability of divorce by 1.4 (fixed-effects) to 3.6 (OLS) to 12.6 (IV) percentage points. These effects represent a change of 7.2-64.9 percent from the mean divorce rate of 19.4 percent. The results for vmen indicate that moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell increase the probability of divorce by 0.11 (OLS) to 1.0 (fixed-effects) to 5.7 (IV) percentage points. These effects represent a change of 0.82-5.7 percent from the mean divorce rate of 13.3 percent.

The instrumental variables estimates in this paper are sizeable, leading one to wonder if they are perhaps too big. There are two reasons to believe that a sizable relation could exist. First, if the workplace is now the primary venue for extra-marital search, then a substantial relationship between occupational sex-mix and divorce is perhaps not so surprising. A very recent book by Shirley Glass, a psychologist and expert in infidelity research, proclaims on page one, "Today's workplace has become the new danger zone of romantic attraction and opportunity."[19] Second, work by Chiappori and Weiss cited above suggests that marriage markets have features that make them highly sensitive to exogenous shocks, such as the infusion of women into the workforce.

If such a sizeable relationship between sex-integration in the workplace and divorce does exist, then it has to be acknowledged that increases in the labor force participation of women do not just cause divorce by raising the incomes of women outside of marriage. There is a second mechanism through which the increased labor force participation of women lowers the costs of extra-marital search.

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[19] Glass, Sherry. 2003. *Not "Just Friends"*. The Free Press: New York.

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Table 1: Distribution of Fraction Female in Occupation and Industry

	5^{th} %ile	25^{th} %ile	Median	75^{th} %ile	95^{th} %ile
Women					
Occupation	0.257	0.499	0.744	0.898	0.989
Industry	0.216	0.477	0.614	0.749	0.892
Industry-Occupation	0.218	0.540	0.800	0.928	0.990
Men					
Occupation	0.020	0.072	0.262	0.404	0.737
Industry	0.106	0.194	0.322	0.530	0.749
Industry-Occupation	0.013	0.053	0.182	0.404	0.741

Notes: Calculations from 1990 PUMS using samples of ever-married, non-widowed non-institutionalized men and women ages 18-55. Those women with industry-occupation cells for which no more than 5 observations are available for calculating the fraction of workers who are female are dropped from the sample. Variables used in table calculations are the fraction of women in each individual's occupation, industry and industry-occupation cell.

Table 2:Fraction Female in Industry-Occupation Cell and Divorce Rates

Fraction Female in Industry-Occupation	Women		Men	
	% of Women in Category	Divorce Rate	% of Men in Category	Divorce Rate
<0.25	5.8%	24.2%	57.8%	13.5%
0.25-0.49	16.4	21.6	25.7	12.5
0.50-0.74	21.6	20.5	11.7	13.7
0.75+	56.2	17.8	4.8	14.7

Notes: Calculations from 1990 PUMS. Sample the same as described in notes of Table 1.

Table 3: Descriptive Statistics

	Women		Men	
	Mean	**St Dev**	**Mean**	**St Dev**
% Divorced	19.4		13.3	
Individual Characteristics:				
Age	37.35	(8.99)	38.72	(8.72)
% Black	7.8		6.4	
% Asian	2.8		2.6	
% Other Race	3.7		4.3	
% Hispanic	1.6		1.7	
% High School Degree	34.0		30.5	
% Some College	32.0		28.0	
% College Degree	14.2		15.4	
% More than College Degree	6.7		9.8	
% Urban	66.1		64.6	
Local PUMA Characteristics:				
Fraction Female	0.51	(0.02)	0.51	(0.02)
Fraction of Men Working	0.92	(0.04)	0.92	(0.04)
Fraction of Women Working	0.78	(0.06)	0.78	(0.06)
	N=1,937,790		N=1,925,662	

Notes: Sample the same as described in notes of Table 1.

Table 4: OLS Estimates of Probability of Divorce, 1990 Census

	Women			Men		
	(1)	(2)	(3)	(4)	(5)	(6)
Fraction Female, Industry-Occupation		-0.0831 (0.0012)	-0.0999 (0.0028)		0.0235 (0.0011)	0.0657 (0.0027)
Fraction Female, Occupation	-0.0476 (0.0012)		0.0356 (0.0026)	0.0160 (0.0013)		-0.0337 (0.0024)
Fraction Female, Industry	-0.0594 (0.0015)		-0.0363 (0.0016)	-0.0018 (0.0015)		-0.0226 (0.0017)
Fraction Female, PUMA	-2.2958 (0.3173)	-2.2866 (0.3173)	-2.2977 (0.3173)	-1.7947 (0.3355)	-1.7699 (0.3355)	-1.7955 (0.3354)
(Fraction Female, PUMA)^2	2.2345 (0.3230)	2.3193 (0.3229)	2.3290 (0.3229)	1.4458 (0.3378	1.4222 (0.3378)	1.4487 (0.3377)
Fraction Men Employed, PUMA	-0.9218 (0.0143)	-0.9162 (0.0143)	-0.9196 (0.0143)	-0.7802 (0.0126	-0.7742 (0.0126)	-0.7762 (0.0126)
Fraction Women Employed, PUMA	0.4350 (0.0093)	0.4363 (0.0093)	0.4354 (0.0093)	0.4648 (0.0080)	0.4612 (0.0080)	0.4627 (0.0080)
		N=1,937,790			N=1,925,662	

Notes: Samples are the same as described in notes of Table 1. Table reports the results from OLS regressions. Dependent variable is a binary indicator for divorce. All regressions include state fixed-effects and individual controls: age, age-squared, race (indicators for black, asian, other), Hispanic origin, urban residence, education (indicators for high school degree, some college, college degree and more than college).

Table 5: OLS Estimates of Probability of Divorce, Wage Controls Added, 1990 Census

		Women			Men	
	(1)	(2)	(3)	(4)	(5)	(6)
Fraction Female, Industry-Occupation		-0.0947 (0.0013)	-0.0974 (0.0030)		0.0091 (0.0012)	0.0266 (0.0029)
Fraction Female, Occupation	-0.0565 (0.0016)		0.0235 (0.0028)	0.0099 (0.0015)		-0.0106 (0.0026)
Fraction Female, Industry	-0.0618 (0.0017)		-0.0321 (0.0018)	-0.0124 (0.0017)		-0.0242 (0.0020)
Mean Male Wage, Industry- Occupation Cell		-0.0012 (0.0001)	-0.0002 (0.0002)		-0.0037 (0.0001)	-0.0029 (0.0002)
Mean Female Wage, Industry-Occupation Cell		-0.0003 (0.0002)	0.0050 (0.0003)		-0.0011 (0.0001)	-0.0004 (0.0002)
Mean Male Wage, Occupation	0.0020 (0.0002)		0.0021 (0.0003)	-0.0040 (0.0002)		-0.0016 (0.0002)
Mean Female Wage, Occupation	-0.0032 (0.0003)		-0.0071 (0.0004)	0.0004 (0.0002)		0.0002 (0.0003)
Mean Male Wage, Industry	-0.0037 (0.0001)		-0.0032 (0.0001)	-0.0016 (0.0001)		0.0000 (0.0002)
Mean Female Wage, Industry	0.0030 (0.0002)		0.0003 (0.0003)	-0.0032 (0.0002)		-0.0030 (0.0003)
Log Male Wage Variance, Industry- Occupation Cell		-0.0022 (0.0005)	-0.0012 (0.0005)		0.0042 (0.0005)	0.0019 (0.0006)
Log Female Wage Variance, Industry-Occupation Cell		-0.0109 (0.0005)	-0.0064 (0.0006)		0.0052 (0.0003)	0.0036 (0.0004)
Log Male Wage Variance, Occupation	-0.0152 (0.0011)		-0.0123 (0.0012)	-0.0047 (0.0010)		-0.0042 (0.0011)
Log Female Wage Variance, Occupation	-0.0172 (0.0012)		-0.0175 (0.0013)	0.0038 (0.0007)		0.0012 (0.0008)

Log Male Wage Variance, Industry	0.0325 (0.0011)		0.0286 (0.0012)	0.0302 (0.0012)		0.0265 (0.0012)
Female Wage Variance, Industry	-0.0434 (0.0015)		-0.0331 (0.0016)	0.0064 (0.0012)		0.0036 (0.0013)
		N=1,907,701			N=1,853,243	

Notes: Samples are the same as described in notes of Table 1, with additional requirement that sufficient observations exist to industry-occupation cell to calculate mean and variance for wages by sex. Table reports the results from OLS regressions. Dependent variable is a binary indicator for divorce. All regressions include local controls: fraction female in PUMA, fraction of men in PUMA working, fraction of women in PUMA working, mean male wage in PUMA, mean female wage in PUMA, state fixed-effects, as well as individual controls: age, age-squared, race (indicators for black, asian, other), Hispanic origin, urban residence, education (indicators for high school degree, some college, college degree and more than college).

Table 6: Fixed-Effects Estimates of Probability of Divorce, Occupation and Industry Fixed Effects, 1990 Census

	Women			Men		
	(1)	(2)	(3)	(4)	(5)	(6)
Fraction Female, Industry-Occupation		-0.0354 (0.0037)	-0.0428 (0.0040)		0.0295 (0.0036)	0.0110 (0.0039)
Fraction Female, Occupation	0.0067 (0.0091)		0.0376 (0.0095)	0.0661 (0.0084)		0.0600 (0.0089)
Fraction Female, Industry	-0.0168 (0.0112)		0.0036 (0.0018)	0.0575 (0.0097)		0.0512 (0.0099)
Mean Male Wage, Industry- Occupation Cell		0.0004 (0.0001)	-0.0001 (0.0002)		-0.0011 (0.0001)	-0.0016 (0.0002)
Mean Female Wage, Industry-Occupation Cell		0.0017 (0.0002)	0.0008 (0.0003)		0.0003 (0.0001)	-0.0002 (0.0002)
Mean Male Wage, Occupation	0.0018 (0.0003)		0.0020 (0.0003)	0.0013 (0.0003)		0.0024 (0.0003)
Mean Female Wage, Occupation	0.0030 (0.0005)		0.0023 (0.0006)	0.0007 (0.0002)		0.0008 (0.0003)
Mean Male Wage, Industry	0.0007 (0.0004)		0.0008 (0.0004)	-0.0023 (0.0004)		-0.0015 (0.0004)
Mean Female Wage, Industry	0.0005 (0.0008)		0.0002 (0.0008)	0.0020 (0.0006)		0.0020 (0.0006)
Log Male Wage Variance, Industry- Occupation Cell		-0.0014 (0.0004)	-0.0002 (0.0005)		0.0007 (0.0005)	0.0014 (0.0006)
Log Female Wage Variance, Industry-Occupation Cell		-0.0022 (0.0005)	-0.0019 (0.0006)		-0.0002 (0.0004)	0.0001 (0.0004)
Log Male Wage Variance, Occupation	-0.0034 (0.0007)		-0.0034 (0.0008)	-0.0015 (0.0008)		-0.0023 (0.0009)

Log Female Wage Variance, Occupation	-0.0014 (0.0008)		-0.0002 (0.0009)	0.0000 (0.0006)		-0.0001 (0.0006)
Log Male Wage Variance, Industry	-0.0023 (0.0011)		-0.0023 (0.0011)	0.0023 (0.0011)		0.0022 (0.0011)
Female Wage Variance, Industry	-0.0012 (0.0010)		-0.0006 (0.0010)	-0.0028 (0.0008)		-0.0027 (0.0008)
		N=1,298,020			N=1,203,736	

Notes: Regressions estimated using a 90% sample of the data sample used in Tables 5 and 6. Regression specifications are the same as those used in Table 5, with the addition of industry and occupation fixed-effects.

Table 7: IV Estimates of Probability of Divorce, 1990 Census

	Women		Men	
	(1)	(2)	(3)	(4)
Fraction Female, Industry-Occupation		-0.3247 (0.0218)		0.1538 (0.0238)
Fraction Female, Occupation	-0.7155 (0.0426)		0.4035 (0.0409)	
Fraction Female, Industry	0.3543 (0.0388)		-0.2087 (0.0348)	
	N=1,907,701		N=1,853,243	

Notes: Samples are the same as described in notes of Table 1. Table reports the results from 2SLS regressions. The occupational and industrial compositions of the PUMA are used as instruments for the fraction female in the occupation, industry and industry-occupation cell as described in the text. All regressions include all the occupation and industry wages controls, PUMA-specific controls, state fixed-effects, and individual-specific controls control variables used in the OLS regressions reported in Table 5.

Table 8: Coefficient on Fraction Female in Industry-Occupation Cell, Age and Race-Specific Samples

	Women		Men	
	(1) OLS	(2) IV	(3) OLS	(4) IV
Full Sample (from Tables 5 and 7)	-0.0948 (0.0012)	-0.4028 (0.0207)	0.0089 (0.0012)	0.2309 (0.0456)
N	1,907,701		1,853,243	
Ages 18-29	-0.0750 (0.0024)	-0.1599 (0.0595)	0.0087 (0.0028)	0.2432 (0.0653)
N	415,499		293,932	
Ages 30-40	-0.1062 (0.0020)	-0.2113 (0.0532)	0.0156 (0.0019)	0.0285 (0.0432)
N	737,606		723,909	
Ages 41-55	-0.0985 (0.0022)	-0.3312 (0.0718)	0.0151 (0.0018)	0.2103 (0.0363)
N	678,184		737,904	
White	-0.1002 (0.0013)	-0.4549 (0.0224)	0.0128 (0.0012)	0.2229 (0.0281)
N	1,627,904		1,593,371	
Non-White	-0.0350 (0.0039)	-0.0007 (0.0413)	-0.0017 (0.0035)	0.1303 (0.0385)
N	259,074		226,337	

Notes: Samples are regression models are the same as those used in columns 2 and 5 of Tables 5 and 7, restricted to the described age or race groups. Sex-mix, wage and wage dispersion measures for each industry, occupation and industry-occupation cell are specific to the age or race group used in the analysis.

Table A1: Summary Statistics for Mean Wage and Wage Dispersion

	Women		Men	
	Mean	**St Dev**	**Mean**	**St Dev**
Mean Wage, Industry-Occupation:				
Male	12.33	(5.00)	14.38	(5.77)
Female	9.29	(3.22)	11.11	(3.71)
Mean Wage, Occupation:				
Male	12.51	(4.38)	14.26	(5.21)
Female	9.78	(2.90)	10.79	(3.06)
Mean Wage, Industry:				
Male	14.11	(5.23)	13.79	(3.87)
Female	9.73	(2.07)	9.99	(1.75)
Mean Wage, PUMA:				
Male	13.31	(3.03)	13.35	(13.35)
Female	9.60	(1.89)	9.61	(9.61)
Wage Variance, Industry-Occupation:				
Male	109.99	(159.71)	124.55	(138.58)
Female	66.73	(77.81)	82.39	(143.54)
Wage Variance, Occupation:				
Male	110.89	(77.07)	128.31	(117.24)
Female	67.44	(36.43)	81.61	(59.77)
Wage Variance, Industry:				
Male	157.75	(134.32)	129.46	(96.35)
Female	70.43	(21.40)	70.32	(21.77)
Wage Variance, PUMA:				
Male	125.28	(72.03)	125.85	(72.96)
Female	68.21	(31.33)	68.39	(31.44)
	N=1,907,701		N=1,853,243	

Notes: Sample the same as described in notes of Table 1.