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

by

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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 and the NLSY79. For the analysis with Census data, two separate strategies are employed to address endogenous occupation and industry choice. In the first, industry and occupation fixed-effects are included in the regression. In the second, the sex-mix a worker faces on the job is instrumented with the industrial and occupational 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 his or her 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 the likelihood of being observed as divorced among ever-married respondents in the 1990 Census and the National Longitudinal Survey of Youth-1979 cohort (NLSY79).

Choice of occupation and industry could be endogenously related to other unobserved characteristics of individuals that make them more or less prone to divorce. In the analysis with Census data, two separate strategies are employed to address this unobserved heterogeneity. In the first, occupation and industry fixed-effects are included in the regression model. Sex-mix measures are calculated at the state-level to facilitate identification in this approach. In the second, the sex-mix 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,

[1] Also cited in Ruggles (1997).

1992; Ruggles, 1997; and South, 2001). These studies do not consider the effect of female employment on workplace contact between men and women. South (2001) discusses 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," 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(e.g. Lerman, 1989; Olsen and Farkas, 1990; Fitzgerald, 1991; Lichter, LeClere and McLaughlin, 1991; Brien, 1998). 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. In contrast, Angrist (2002) uses exogenous variation in immigration flows to study the effects of sex ratios within immigrant groups on marriage outcomes of first and second-generation immigrants.

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. Aberg (2003) is the only study closely related to the analysis conducted in this paper. Using data on Swedish firms, she also finds divorce rates are higher for married workers in cases where a large fraction of coworkers are of the opposite sex. Specifically, she finds that a person is 70 percent more likely to divorce if 100 percent of co-workers are of the opposite sex and of similar age, compared to if they are all of the same sex or considerably younger or older than the individual. In her data, she has the benefit of observing the sex and marital status of the co-workers in an individual's firm, as opposed to the sex-mix of workers in an individual's occupation or industry. On the other hand, she does not address endogenous choice of industry and occupation in her analysis. Nor does she have data on spouse's occupation, as is available in the NLSY79.

The theoretical literature suggests that sex-integration in the workplace should 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 is 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.[2]

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 for individuals who are married, search for alternative mates outside of the workplace is extremely limited and very costly compared to that for single

[2] 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.

individuals.[3] So while sexual-integration of the workforce lowers search costs for singles, which should result in higher-quality matches that are less prone to divorce, it is likely that this effect is dominated by the increased ability of married individual's to search among alternative mates. This ultimately is an empirical question.

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 does not produce another 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 meets many 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.[4] 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

[3] Lauman et al (1994) report that about 15 percent of individuals met their spouse through work.
[4] A similar point is made by South and Lloyd (1995).

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.

3. Empirical Analysis with 1990 Census Data

A. Sex-Segregation by Occupation and Industry

Other studies have documented that male and female workers are heavily segregated by occupation and, to a lesser extent, by industry. This feature of the labor market has 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 the 52,709 observed civilian industry-occupation combinations. The statistic of interest is the fraction of workers between the ages of 18 and 55 who are female. Distributions of these sex-mix statistics for the sample used in the regression analysis are reported in Table 1.[5] It is clear that there is still substantial segregation by industry and occupation. The median woman in the regression sample works in an

[5] As described below, the sample from the 1990 PUMS used in the regression analysis is all ever-married, non-widowed, non-institutionalized individuals ages 18-55 that report an industry and occupation. Very small industry-occupation cells are omitted from the sample as described below. To be clear, the sex-mix measures are calculated using *all* workers ages 18-55 in industry, occupation or industry-occupation cell, as opposed to only the ever-married workers used in the regression sample.

occupation that is 74% female and an industry that is 61% female, while the median man in the regression sample works in an occupation that is 27% 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 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 clear relationship between percent female and divorce. Only 5.7 percent of women work in industry-occupation combinations that are less than 25% female, but their divorce rate is 24.2 percent. In contrast, 55.6% 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.

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 18 to 55 who report an industry and occupation.[6] Individuals are dropped from the sample if their industry-occupation cell is too small to calculate the sex-mix measure and wage controls used in the regression analysis. Specifically, industry-occupation cells with no more than 5 observations or without two male

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

workers and two female workers with wages in the range of $2-$200/hour are omitted from the sample.[7] The final sample consists of 1,907,701 women and 1,853,243 men.

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. Among non-institutionalized ever-married women ages 18-55 in the 1990 PUMS, 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 Regression Model

The baseline regression model used is the linear probability model:

$$\begin{aligned} Y_{ionps} = {} & \beta_0 + \beta_1 FractionFemale\_INDOCC_{on} + WageControls\_INDOCC_{on}\beta_2 \\ & + \beta_3 FractionFemale\_PUMA_p + \beta_4 (FractionFemale\_PUMA_p)^2 + PUMAControls_p \beta_5 \\ & + IndividualControls_{ionps}\beta_6 + STATE_s \delta + (STATE_s * Urban_i)\phi + \varepsilon_{ionps} \end{aligned} \tag{1}$$

Where for person *i* working in occupation *o* and industry *n*, living in PUMA *p* in state *s*, *Y* is an indicator for divorce. *FractionFemale_INDOCC* is, for the worker's industry-occupation cell, the fraction of workers ages 18-55 who are female. To be clear, the fraction female in industry-occupation cell is calculated using *all* workers ages 18-55. The regression sample is then further restricted to ever-married workers. *WageControls_INDOCC* is a vector of wage controls for the worker's industry-occupation cell. Specifically, mean male and female wages and the logarithms

[7] Dropping industry-occupation cells with 5 or few workers omits 39,991 individuals, a little less than 1 percent of the sample. Additionally, only wage observations in the rage of $2-$200/hour are used in the calculation of the wage controls. Two workers of each sex with wage observations in this range are necessary to calculate the wage dispersion measures for an industry-occupation cell. Dropping industry-occupation cells without two wage observations for each sex omits another 102,499 workers, or another 2.7% of the sample**.**

of male and female wage variances are calculated from the 1990 PUMS for each industry-occupation cell. *FractionFemale_PUMA* is the fraction of residents of the PUMA ages 18-55 who are female. *PUMAControls* is a vector of PUMA-specific economic controls that includes the fraction of men employed in the PUMA, the fraction of women employed in the PUMA, the mean male and female wages in the PUMA and the logarithms of male and female wage variances for the PUMA. *IndividualControls* is a vector of individual control variables, which includes age, age-squared, race indicators (black, asian, other), a Hispanic ethnicity indicator, an urban residence indicator, and education indicators (high school degree, some college, college degree, more than college degree). *STATE* is a vector of state indicator variables and *STATE*Urban* interacts the state indicators with an indicator for urban residence. These state fixed-effects and state-urban fixed-effects control for unobserved differences across states and differences between urban and rural areas within states. Descriptive statistics for variables other than the wage controls are reported in Table 3. Descriptive statistics for the wage measures are reported in Appendix Table A1.

The cross-sectional nature of the data limits our information regarding marital status. The Census only identifies whether an individual is married, divorced, separated, widowed or has never married at the time of the survey. If an individual has previously divorced and then remarried, we only observe them as currently married, we do not know that they were previously divorced. In this paper, those reported as married, divorced or separated are included in the sample and only those that report they are currently divorced are categorized as such. The analysis with the Census will not capture the effect of workplace contact that generate divorces that are quickly followed by remarriage. But to the extent that workplace contact, through the mechanisms discussed above, generates divorce that is not quickly followed by remarriage, part of the effect of interest can be identified in the cross-sectional census data.[8]

One further concern might be that the sex-mix in an individual's workplace is likely different from the sex-mix in that worker's industry-occupation cell. If the form of the measurement error is classical, so that the workplace sex-mix is merely the industry-occupation sex-mix plus some random error, the OLS estimates will not suffer from attenuation bias. This is because the industry-occupation sex-mix is an aggregation of workplace sex-mix across the industry-occupation cell and, as a result, all variation in industry-occupation sex-mix is "true"variation rather than white noise. The industry-occupation sex-mix is less noisy than the workplace sex-mix.[9]

D. OLS Regression Results

The initial regression results are reported in Table 4. Columns 1 and 2 report the results for women and columns 3 and 4 report the results for men. Columns 1 and 3 report the results obtained from estimating the regression model specified in equation (1). For women, working in an industry-occupation with a higher fraction of female workers lowers the probability of divorce. For men, working in an industry-occupation with a higher fraction female raises the probability of divorce. In columns 2 and 4, fraction female in occupation and fraction female in industry replaces the fraction female in the industry-occupation cell. The wage controls for these models are the mean male and female wages for industry and occupation and the logarithms of

[8] 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).

[9] In fact, if we had data on workplace sex-mix that we believed to be measured with error, fraction female in industry-occupation cell would be a valid instrument for workplace sex-mix under the classical measurement error model.

male and female wage variances for industry and occupation. 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.[10]

The magnitude of the effect for men is quite 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 .534 to .924 would decrease her probability of divorce by 3.7 percentage points. A man moving from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .059 to .415, would increase his probability of divorce by 0.32 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.

The coefficient estimates for the occupation, industry and industry-occupation wage controls are also reported in Table 4. 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

[10] A model that simultaneously measures the effects of fraction female in industry, occupation and industry-occupation cell is difficult to interpret. Estimates from specifications of this type typically produced coefficients on the fraction female in the industry-occupation cell that were consistent with expectations (negative for women and positive for men), but the fraction female in industry and fraction female in occupation frequently had a counter-intuitive sign. 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.

occupation or industry with above-average wages, this suggests that his earnings potential is also above average.[11] 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. The wage results are therefore not a primary focus of the paper. The general finding in Table 4 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 are exceptions in both cases.

The other coefficients reported in Table 4 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 and higher wages for men in the local area are associated with a lower probability of divorce and a higher employment rate and higher wages for women in the local area are associated with a higher probability of divorce.

E. 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

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

ages, which should increase the quality of the match.[12] 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 solution is to include industry and occupation-specific fixed-effects in the model. This will sweep out any unobserved industry and occupation characteristics. For example, if women in more male-dominated occupations work longer hours and have fewer children than women in more traditional occupations, these fixed-effects will purge out mean fertility and mean hours of work by occupation and industry. The regression specification with industry-occupation controls used in columns 1 and 3 of Table 4 is identified even with the inclusion of industry and occupation fixed-effects. The specification with separate industry and occupation controls used in columns 2 and 4 of Table 4 is not identified with industry and occupation fixed-effects in the model. In this case, the model can be estimated by first calculating the sex-mix measures for industry and occupation at the state, rather than national level. This provides the additional variation in the sex-mix measures necessary for identification.[13]

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

[12] For the regression models estimated in Table 4, 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 5.6-9.0 percentage point lower probability of divorce than those without a high school diploma. For men, the correspondent results are .35-2.2 percentage points lower probability of divorce for those with a high school diploma but no college degree and 4.9-5.4 percentage points lower probability of divorce for those with at least a college degree.

[13] Because of the requirement that there be more than five workers to calculate sex-mix and wage statistics for industry-occupation cell, calculating these measures at the state level reduces the sample by 28% to 1,337,791 for men and by 24% to 1,442,226 for women. Estimating the models in Table 4 on these reduced samples produces estimates similar to those reported in Table 4.

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. 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. Therefore, the residual variation available to identify the effect of interest is greatest for models using sex-mix in industry-occupation cell and is greater in the male sample than the female sample.

The results of the fixed-effects analysis are reported in Table 5.[14] The results for women are reported in columns 1-3 while the results for men are reported in columns 4-6. Columns 1 and 4 report the results from adding occupation and industry fixed-effects to the basic model estimated in columns 1 and 3 of Table 4. For women, the effect of fraction female in industry-occupation cell is negative and about 35% smaller in magnitude than that estimated in Table 4. For men, the effect of fraction female in industry-occupation is positive and about 50% larger in magnitude than that estimated in Table 4. Columns 2 and 5 report the results of estimating the same model as in columns 1 and 3, but with the sex-mix and wage controls for each industry-occupation cell calculated at the state, rather than national, level. The coefficient on fraction female in industry-occupation is negative for women and positive for men, and unlike the results in Table 4, the magnitudes are relatively similar for men and women. Columns 3 and 6 report results based on the model used in columns 2 and 4 of Table 4, with the addition of industry and occupation fixed-effects and with the industry and occupation-specific sex-mix measures and wage controls calculated at the state level. For women, sex-mix in industry and occupation has

[14] Due to computer memory constraints, the results in columns 1 and 4 of Table 5 are estimated using an 80% sample of the 1,907,701 observations available for women and the 1,853,243 observations available for men. Due to the same constrains, the results in columns 2-3 and 5-6 of Table 6 are estimated using a 90% sample of the 1,442,226 observations available for women and the 1,337,791 observations available for men. Footnote 12 above details why there are fewer observations available when the sex-mix measures are calculated at the state level.

no effect on divorce. For men, working in occupations and industries with more women increases the probability of divorce. Overall, the sex-mix measures for which there are the most available variation independent of the fixed effects, those for men and those calculated at the state level by industry-occupation cell, are the ones for which the coefficient estimates 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 25^{th} percentile to the 75^{th} 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. The coefficient estimate in column 5 indicates that moving a man from the 25^{th} percentile to the 75^{th} percentile of fraction female in industry-occupation cell would increase his probability of divorce by 1.1 percentage points, or 7.9% 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.

F. IV Analysis

As an alternative approach to address 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. This instrument varies by sex and by PUMA, but not by industry or occupation. For a male worker in PUMA *p*, the instrument for the fraction employment in a worker's occupation that is female is:

$$IVMaleOCC_p = \sum_o ShareMaleEmp_{op} * FractionFemale\_OCC_o \, , \qquad (2)$$

where $ShareMaleEmp_{op}$ is the fraction of total male employment in PUMA *p* that occurs in occupation *o* and $FractionFemale\_OCC_o$ is the fraction of *national* employment in occupation *o* that is female. An analogous instrument can be calculated for the fraction female in a male worker's industry:

$$IVMaleIND_p = \sum_n ShareMaleEmp_{np} * FractionFemale\_IND_n \text{ ,} \tag{3}$$

where $ShareMaleEmp_{np}$ is the fraction of total male employment in PUMA *p* that occurs in industry *n* and $FractionFemale\_IND_n$ is the fraction of national employment in industry *n* that is female.

The instruments for a female worker in PUMA *p* are:

$$IVFemOCC_p = \sum_o ShareFemEmp_{op} * FractionFemale\_OCC_o \text{ ,} \tag{4}$$

and:

$$IVFemIND_p = \sum_n ShareFemEmp_{np} * FractionFemale\_IND_n \text{ .} \tag{5}$$

These instruments are calculated for each of the 1725 PUMAs in the 1990 PUMS.

The instruments are therefore weighted averages of the industry and occupation-specific sex-mix measures, where the weights are the shares of local male and female employment in those industries and occupations. Each instrument is therefore an expected value for fraction female in industry or occupation given a worker's sex and PUMA of residence. A worker living in an area where most employment is in occupations and industries that are typically highly segregated by sex is likely to experience relatively little sex integration. The reverse is true for a worker living in an area where more employment is in occupations and industries that tend to be integrated by sex.

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, but this IV approach does. It could be, 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 there 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 6. Columns 1 and 2 instrument the sex-mix measures for female workers with the variables described in equations (4) and (5). Columns 3 and 4 instrument sex-mix measures for male workers with the variables described in equations (2) and (3).

In columns 1 and 3, the fraction female in the worker's industry-occupation cell is instrumented with both the occupational and industrial composition variables.[15] In both cases the

[15]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

results are of the predicted sign and the effects are larger in magnitude than those obtained with OLS or fixed-effects estimation. These estimates indicate that moving a woman from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .534 to .924 decreases her probability of divorce by 15.7 percentage points. Moving a man from the 25th percentile to the 75th percentile of fraction female in industry-occupation cell, from .059 to .415, would increases his probability of divorce by 8.4 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.

In column 2, the fraction female in a woman's occupation, as predicted by the occupational composition of the female workforce in her PUMA, has the predicted negative effect on divorce and is again much larger than the estimates reported in Tables 4 and 5. 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, the results in column 4 display a similar pattern. The predicted fraction female in their occupation has the expected positive effect on divorce and is large in magnitude. The predicted fraction female in the industry, however, has a negative, although smaller, effect.

G. Age and Race Specific Results

We might expect the effect of sex-mix on divorce to vary by age and race. Table 7 reports OLS, fixed-effects 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 1 and 3 of Table 4, columns 1 and 4 of Table 5 and columns 1 and 3 of Table 6. The second row reports the results for women and men ages

---

industry-occupation cell. For women, the first-stage partial F-statistic on the instruments described in equations (4) and (5) is 3220 where the first-stage partial F-statistic on an instrument using employment in industry-occupation cells is 621. For men, the analogous F-statistics are 2145 and 942.

18-29. The sex-mix measures, the wage and wage dispersion measures, and the PUMA-specific instruments have all been re-calculated on this sample of young adults so that these variables are specific to 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 remained small.

4. NLSY Analysis

This section extends the analysis in this paper to data from the NLSY79, a panel data set based on annual surveys of men and women who were 14-21 years old on January 1, 1979. Respondents were first interviewed in 1979, re-interviewed each year through 1994, and have been interviewed every two years since 1994. The analysis in this section uses data from 1979-2000. The NLSY79 affords two primary advantages over the Census data. First, the NLSY79 contains longitudinal data with marital histories. Unlike the cross-sectional marital status information in the Census, I do not fail to observe a divorce because an individual remarries. First marriages can be separated from later marriages. The second primary advantage of the NLSY79 data is that there is information on the occupation of the spouse. Therefore, it is possible to estimate the effect of occupational sex-mix of both members of the couple on divorce. An additional advantage of the NLSY79 data is that it provides a richer set of individual characteristics to use as controls.

These advantages come at a cost. The primary disadvantage of the NLSY79 data compared to the 1990 Census is the substantial reduction in sample size. This will limit the potential to use the fixed-effects and instrumental variables strategies employed above to deal with endogenous choice of industry and occupation. An additional disadvantage is that the NLSY79 is a relatively young sample, with respondents ranging in age from 35 to 42 in 2000, the last year of data used in this analysis. This limits the number of divorces we observe in the data, and particularly the number of divorces occurring in marriages of long duration.

Because the primary advantages of the NLSY79 data can be exploited using simple cross-sectional analysis, the first set of results is obtained using linear probability models similar in specification to those used to analyze the 1990 Census data. Further analysis then exploits the longitudinal nature of the data using discrete-time hazard models.

A. OLS Analysis

Initial OLS analysis with the NLSY79 data is formulated to be conceptually similar to the cross-sectional analysis with the Census data. The analysis sample consists of all ever-married respondents reporting the necessary industry and occupation information.[16] Only first marriages are used in the analysis. The respondent's industry and occupation reported for the year of marriage are used to calculate the respondent's sex-mix in industry-occupation cell. Spouse's occupation reported for the year of marriage is used to calculate the sex-mix in the spouse's occupation. Spouse's industry is not reported in the NLSY79 data. The industry and occupation information from the year of marriage is used to calculate a single cross-sectional measure of job sex-mix in the hopes that choice of industry and occupation at the beginning of the marriage is

[16] The military over-sample is excluded from analysis.

more exogenous to the stability and quality of the marriage than industry and occupational choice in later years. For respondents with missing industry or occupation information for the year of marriage, information from most recent job reported in the past 5 years is used. Because spousal information is not reported prior to the year of marriage, occupation of spouse cannot be filled in from prior information.

The regression model is a linear probability model of the form:

$$\begin{aligned} Y_{iondps} = {} & \beta_0 + \beta_1 FractionFemale\_INDOCC_{on} + \beta_2 FractionFemale\_SpouseOcc_d \\ & + WageControls_{ond}\beta_3 + LocalControls_p\beta_4 + IndividualControls_i\beta_5 \\ & + STATE_s\delta + (STATE_s * Urban_i)\phi + OCC_o\gamma_1 + IND_n\gamma_2 + SpouseOCC_d\gamma_3 + \varepsilon_i \end{aligned} \tag{6}$$

Where for person *i* working in occupation *o* and industry *n*, with a spouse working in occupation *d*, living in local area *p* in state *s*, Y is an indicator that equals one if the individual reports ending their first marriage in divorce at any time in the NLSY survey. *FractionFemale_INDOCC* is, for the respondent's industry-occupation cell at the time of marriage, the fraction of workers ages 18-55 who are female. *FractionFemale_SpouseOcc* is, for the spouse's occupation at the time of marriage, the fraction of workers ages 18-55 who are female. *WageControls* is a vector containing mean male and female wages for respondent's industry-occupation cell and mean male and female wages for spouse's occupation.[17] *LocalControls* includes characteristics of the respondent's country group as calculated from the 1980 Census for marriages from 1979-1985 and for characteristics of the respondent's PUMA as calculated from the 1990 Census for marriages after 1985. The local area characteristics include the fraction of local residents ages 18-55 who are female and its square, the fraction of men employed in the local area, the fraction of women employed in the local area, and mean male and female wages in the local area. *IndividualControls* is a vector of individual control variables, described below. *STATE* is a

[17] For marriages in years 1979-85, the sex-mix and wage measures are calculated from 1980 Census data. Sex mix and wage measures calculated from 1990 Census data are used for marriages after 1985.

vector of state indicator variables and *STATE*Urban* interacts the state indicators with an indicator for urban residence. The local controls, state fixed-effects and state-urban effects are all based on location at the time of marriage. *OCC* is a vector of occupation fixed-effects, *IND* is a vector of industry fixed-effects, and *SpouseOCC* is a vector of spouse occupation fixed-effects, all calculated at the 1-digit or 2-digit code level. Because the NLSY79 is a stratified sample, the regression is weighted using the initial weights reported for the 1979 survey.

The individual controls used in the OLS analysis are the age of first marriage, race/ethnicity indicators (black, Hispanic), highest grade completed, highest grade completed of spouse, indicator for living with both biological parents at age 14, the respondent's expected age of marriage (measured in 1979), and expected number of children (measured in 1979).[18] Table 8 reports descriptive statistics for the NLSY sample.[19] While a total of 8,553 first marriages are reported in the NLSY79, 3,047 of which end in divorce during the survey, only 5,109 marriages are observed with the necessary industry, occupation and spouse occupation information to be included in the analysis. 32 percent of the excluded marriages are missing the necessary occupation and industry information because they occur before 1979, the first year of the survey. Many of the remaining marriages with missing data on industry and occupation are marriages by young respondents, who marry before they ever work. Of the 5,109 marriages with the necessary industry and occupation information, 1,578, or roughly 30 percent, of these marriages end in divorce during the survey. The statistics reported in Table 8 indicate that because men tend to marry at later ages than women, the marriages reported by male respondents tend to

[18] Expected age of marriage is a categorical variable: 1:<20, 2:20-24, 3:25-29, 4:30+, 5: Never.
[19] Because Table 8 reports unweighted means and the NLSY79 over-samples non-whites, black and Hispanic respondents make up over 30 percent of the sample. Weighted means, using 1979 weights, would show them as 12-13 percent of the sample. None of the other statistics reported in Table 8 change appreciably when weights are used.

occur slightly later in the survey. As a result, the divorces reported by male respondents tend to occur earlier in the marriage.

Estimates were also obtained using an expanded set of controls. These additional controls included the age difference between the respondent and spouse, indicator for foreign birth, indicator for living in the South at age 14, indicator for urban residence at age 14, indicator for living with a single mom at age 14, mother's completed years of education, father's completed years of education, indicator for Protestant upbringing, indicator for Catholic upbringing, indicator for upbringing in another religion, birth year fixed-effects, interactions of birth year fixed-effects with expected age of marriage, interactions of birth year fixed-effects with expected number of children, and the logarithm of male and female wage variances for industry-occupation cell, spouse's occupation and local area. Because adding the additional set of controls decreased the sample size by 20 percent, the additional controls were rarely statistically significant, and the coefficient estimates on the sex-mix measures changed relatively little with the additional controls, the results reported here are estimated using the smaller set of controls.

The results obtained from estimating equation (6) are reported in Table 9.[20] The first two columns report the results for women, with the first column using 1-digit industry and occupation code fixed-effects and the second column using 2-digit industry and occupation code fixed-effects.[21] The results are consistent with expectations. For a married woman, working in an industry-occupation with a higher fraction of women lowers her probability of divorce. A higher fraction of women in her husband's occupation increases her probability of divorce. The

[20] The sample sizes in Table 9 are smaller than those in Table 8 due to missing data on additional controls.
[21] There are nine 1-digit occupation code fixed-effects and ten 1-digit industry code fixed-effects, as well as nine 1-digit spouse's occupation code fixed-effects. There are 84 2-digit occupation code fixed-effects and 88 2-digit occupation code fixed-effects, as well as 84 spouse's occupation code fixed-effects.

coefficient on fraction female in the woman's industry-occupation cell is strongly significant using both 1-digit and 2-digit fixed-effects. The coefficient on fraction female in husband's occupation is only significant using 1-digit fixed-effects. In both cases, the effects are sizeable. Using the metric of moving from the 25th percentile to the 75th percentile of the sex-mix measure as reported in Table 1, a woman moving from an industry-occupation sex-mix measure of .534 to .924 would decrease her probability of divorce by 6.8-9.5 percentage points. Her husband moving from an occupational sex-mix of .072 to .405 would increase her probability of divorce by 3.8-5.1 percentage points.

The results for male respondents are reported in columns 3 and 4 of Table 9. While the coefficient on fraction female in wife's occupation has the expected negative sign in both cases, the coefficient on the fraction female in the man's industry-occupation cell only becomes positive when 2-digit fixed-effects are used. The coefficient estimates are all small in magnitude and statistically insignificant. If we interpret the magnitudes of the coefficient estimates in column 4 despite the lack of statistical significance, they suggest that moving the male respondent from the 25th to 75th percentile of the relevant sex-mix measure increases the probability of divorce by 1.7 percentage points. Moving his wife from the 25th to the 75th percentile decreases the probability of divorce by 2.8 percentage points. These are relatively small effects given that almost a third of the sample divorces during the survey.[22] [23]

B. Discrete-Time Hazard Model Analysis

[22] Instrumental variable analysis using the instruments described in equations (2)-(5) was attempted. The first-stage partial F-statistics were almost uniformly less than 10. The instruments are too weak in this small a sample to perform well.

[23] Additional sensitivity analysis was performed for Table 9. Results obtained restricting the sample to white respondents are consistent with those reported here, with the coefficient estimates slightly larger in magnitude. If industry and occupation from age 25 are used rather than from year of marriage, the results again display the same patterns, although the results for women are smaller in magnitude and the coefficient on sex-mix in industry-occupation is negative in both of the male regressions. If observations currently excluded from the sample due to missing data on industry, occupation or spouse's occupation are included using the first job reported for the respondent or spouse after the year of marriage, the results are similar but the coefficients are smaller in magnitude.

The analysis in this section uses a discrete-time hazard model of the form:

$$
\begin{aligned}
H(t) &= \Pr(\text{Divorce in Year } t| \text{ Married but Not Divorced in Year } t\text{-}1) \\
&= F[\beta_0 + \beta_1 FractionFemale\_INDOCC_{ont} + \beta_2 FractionFemale\_SpouseOcc_{dt} \\
&\quad + WageControls_{ondt}\beta_3 + LocalControls_{pt}\beta_4 \\
&\quad + IndividualControls_i\beta_5 + STATE_{st}\delta + (STATE_{st} * Urban_{it})\phi \\
&\quad + OCC_{ot}\gamma_1 + IND_{nt}\gamma_2 + SpouseOcc_{dt}\gamma_3 + Year_t\lambda + g(Year_t - Year\ of\ \ Marriage_i)]
\end{aligned} \tag{7}
$$

for person *i* working in occupation *o* and industry *n*, living in local area *p* in state *s* in year *t*. In this specification, the sex-mix measures are calculated for industry, occupation and spouse's occupation at time *t*. Similarly, wage controls, local controls, state fixed-effects, state-urban fixed-effects, and industry and occupation code fixed-effects are all measured based on location, industry, occupation and spouse's occupation at time *t*. The individual controls are the same as those used in Table 9. Industry, occupation and spouse's occupation code fixed-effects are at the 1-digit level. Year effects are included in the model and the baseline hazard, g(.) is a vector of dummy variables for duration of marriage, where the hazard is assumed to be constant after 10 years of marriage.[24] Assuming that F[.] is logistic, this specification amounts to estimating logit models where each observation represents a year of marriage for a respondent. Initial 1979 weights are used to weight the regressions.[25]

Using the hazard model expands the sample used for analysis. Individuals who married prior to 1979 can now be included in the analysis, as long as they do not divorce prior to 1979.[26] Individuals who do not report an industry or occupation the year they get married or prior to getting married can now be included in the sample for the years in which industry and

[24] There is right-censoring for cases in which the marriage does not end by 2000 or the individual drops out of the survey prior to 2000. Respondents are assumed to be right-censored in year X if they do not interview from year X to 2000.

[25] NLSY79 also reports updated weights for each year to account for attrition. Results using these annual weights are very similar to those reported in the paper.

[26] There are 7466 first marriages marriages that do not end in divorce prior to 1979. 2,413 of these marriages end in divorce during the survey.

occupation are reported. The same is true for cases in which spouse's occupation is not reported in the initial year of marriage.

If the hazard model is estimated only using observations in which industry, occupation and spouse's occupation are reported for that year, this will generate a sample that is heavily selected on labor force participation. This is problematic given that labor supply is endogenous to marriage and divorce decisions. Two alternative approaches are used to better deal with non-participation. The first is very similar to the occupation and industry measures used in the Census. If occupation, industry or spouse's occupation is missing in a given year, information from the most recent job reported in the past 5 years is used. If there is no job information within the past 5 years, then the observation drops from the sample. This approach produces a sample substantially less selected on labor force attachment. The drawback of this approach is that the sex-mix measures will sometimes reflect the sex-mix the respondent faced on a job they held many years ago, which if they are not currently working, is less likely to generate a divorce.

The second specification directly controls for employment in a given year, understanding that labor supply is endogenous to marital status. If an individual reports positive weeks of work in a given year, then the sex-mix measure is sex-mix for the most occupation and industry reported within the past 5 years. If the individual does not report positive weeks of work in a given year, then the sex-mix measure is set to zero. This is effectively an interaction between an employment indicator and the sex-mix an individual would experience if they chose to work. The same procedure is used for the sex-mix in spouse's occupation. These two sex-mix measures are then used in the analysis while also controlling for employment of both the respondent and the spouse.[27]

[27] Initial analysis of the data revealed that respondents were substantially less likely to report employment for their spouses the year of their divorce. While some of this decline in employment is probably real and is a causal factor

The results of the hazard model analysis are reported in Table 10. The top panel reports the results obtained using sex-mix in most recent job in the past 5 years. Once again, the results for women conform to expectations. For a female respondent, a higher fraction of female workers in her industry-occupation cell reduces the probability of divorce, while a higher fraction of female workers in her spouse's occupation increases the probability of divorce. The coefficient for fraction female in industry-occupation is strongly statistically significant, but the coefficient for fraction female in spouse's occupation is not significant.[28] For men, the results are similar to those in Table 9; both coefficients are small, negative and insignificant.

The magnitudes of the effects can be calculated by setting all other controls to their sample means and calculating the effect of the inter-quartile move on the one-year divorce probability. For a woman with average characteristics, moving from the 25th to 75th percentile of the sex-mix measure (0.534 to 0.924) decreases her 1-year divorce probability from 1.7 percent to 1.26 percent, a decrease of 26 percent. For a woman with average characteristics, moving her husband from the 25th to the 75th percentile of the sex-mix measure (0.072 to 0.405) increases her 1-year divorce probability from 1.38 to 1.55, an increase of 11 percent.

The bottom panel of Table 10 reports the results from the alternative specification that controls for employment. For women, a higher fraction of female workers in industry-occupation cell reduces the probability of divorce and the effect is statistically significant. A higher fraction female in husband's occupation increases the probability of divorce, although the coefficient is very small in magnitude and insignificant. Also as expected, the woman's employment is associated with a higher probability of divorce while the husband's employment

in the divorce, the very low spousal employment rates for the year of divorce suggested that many respondents do not know or refuse to report their spouse's employment status the year their divorce is finalized. Therefore, for cases in which the respondent's spouse worked the year prior to divorce, it was assumed they worked the year of divorce as well.

[28] Standard errors are robust standard errors clustered by respondent.

is associated with a lower probability of divorce, although both coefficients are statistically insignificant. The estimates for women in the bottom panel of Table 10 indicate that moving a woman with average characteristics from the 25th to 75th percentile of the sex-mix measure decreases her 1-year divorce probability from 1.64 percent to 1.35 percent, a decrease of 18 percent.

The results for men are reported in the second column. The fraction female in wife's occupation has the expected negative sign, although it is insignificant. Once again, the fraction female in the man's own industry-occupation cell is also negative and insignificant. The labor supply results conform to expectations in sign, with the man's employment lowering the probability of divorce and his wife's employment associated with a higher probability of divorce, although these coefficients are statistically insignificant as well.[29]

The lack of agreement between the results for female respondents and the results for male respondents is puzzling. While it is not clear why these differences exist, it is at least possible to discuss some of the differences in the data for male and female respondents that might contribute to this lack of symmetry. The first difference between the male and female sample, as discussed in Table 8, is that the men typically marry later in the survey so that there are fewer observations on marriage, and fewer marriages of long duration, in the male sample. The second difference is that because the respondent reports all information for both herself and her spouse, the employment information for both the husband and wife is reported by the wife in the female sample and by the husband in the male sample. Therefore, if there are differences in the

[29] The analysis in Table 10 was also replicated with the substantially expanded set of individual controls described above for Table 9. The results with the larger set of controls largely conform to those reported here, the one exception being that the coefficient on employment for men inexplicably becomes positive in the second column, suggesting that men who work are more likely to divorce. The coefficient, however, remains insignificant. Additional sensitivity analysis of the results in Table 10 included restricting the sample to white respondents and using 2-digit industry and occupation code fixed-effects. The results obtained for white respondents are similar to those reported here, with effects of somewhat larger magnitude. The results obtained using 2-digit code fixed-effects are consistent with those obtained here, but all of the coefficient estimates are statistically insignificant.

accuracy with which men and women report their own and their spouse's employment information, this will generate differences between the two samples.[30] Finally, because there is relatively little background information on the spouse, the individual controls are largely for the wife in the female sample and are largely for the husband in the male sample.

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 from the analysis of 1990 Census data indicate that moving a woman from the 25^{th} percentile to the 75^{th} percentile of fraction female in industry-occupation cell decrease the probability of divorce by 1.4 (fixed-effects) to 3.7 (OLS) to 15.7 (IV) percentage points. These effects represent a change of 7.2-80.9 percent from the mean divorce rate of 19.4 percent. The Census results also indicate that moving a man from the 25^{th} percentile to the 75^{th} percentile of fraction female in industry-occupation cell increase the probability of divorce by 0.32 (OLS) to 1.1 (fixed-effects) to 8.4 (IV) percentage points. These effects represent a change of 2.4-63.2 percent from the mean divorce rate of 13.3 percent.

The results from the analysis of the female sample from the NLSY79 indicate that moving a woman from the 25^{th} to 75^{th} percentile of fraction female in industry-occupation cell

[30] Female respondents in the analysis sample used in Table 10 report an employment rate of 77 percent for themselves and 81 percent for their spouses. Male respondents in the same sample report an employment rate of 90 percent for themselves and 63 percent for their spouses.

decreases her probability of divorce by 6.8-9.5 percentage points and, for the average woman in the NLSY79, decreases her one-year divorce probability by 18-26 percent. Moving her spouse from the 25th to 75th percentile of fraction female in occupation increases his probability of divorce by 3.8-5.1 percentage points and increases his 1-year divorce probability as much as 11 percent. The results for the male sample from the NLSY79 are statistically insignificant in all cases.

Some of the estimates in this paper are sizeable, leading one to wonder if they are perhaps too big. There are three reasons to believe that a sizable relation does exist. First, if the workplace is now the primary venue for extra-marital search, a substantial relationship between occupational sex-mix and divorce is perhaps not so surprising. A 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.”[31] Second, work by Chiappori and Weiss (2001) discussed above suggests that marriage markets have features that make them highly sensitive to exogenous shocks, such as the infusion of women into the workforce. Finally, the large effects obtained in this analysis are consistent with those found using data on Swedish firms by Aberg (2003).

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 in which the increased labor force participation of women lowers the costs of extra-marital search.

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[31] Glass, Shirley. 2003. *Not “Just Friends”*. The Free Press: New York.

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Table 1: Distribution of Fraction Female in Occupation, Industry and Industry-Occupation

	5^{th} %ile	25^{th} %ile	Median	75^{th} %ile	95^{th} %ile
Women					
Occupation	0.257	0.492	0.738	0.898	0.989
Industry	0.222	0.477	0.614	0.749	0.892
Industry-Occupation	0.219	0.534	0.800	0.924	0.989
Men					
Occupation	0.020	0.072	0.270	0.405	0.737
Industry	0.106	0.194	0.324	0.538	0.749
Industry-Occupation	0.016	0.059	0.192	0.415	0.750

Notes: Calculations from 1990 PUMS. Sex-mix measures are the fraction of workers ages 18-55 that are female by occupation, industry and industry-occupation cell. Distributional statistics are calculated for the sample used for regression analysis: ever-married, non-widowed, non-institutionalized men and women ages 18-55 in the 1990 PUMS. Industry-occupation cells in which there are fewer than 5 observations overall and fewer than 2 wage observations each for men and women in the range of $2-$200/hr are dropped from the sample. There are 1,907,701 women and 1,853,243 men in the regression sample.

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	5.7%	24.2%	56.2%	13.5%
0.25-0.49	16.7	21.6	26.7	12.5
0.50-0.74	22.0	20.5	12.1	13.6
0.75+	55.6	17.8	5.0	14.7

Notes: Calculations from 1990 PUMS. Sample is described in notes of Table 1.

Table 3: Descriptive Statistics, 1990 Census

	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		2.7	
% Other Race	3.7		4.3	
% Hispanic	1.7		1.7	
% High School Degree	33.9		30.1	
% Some College	31.9		27.9	
% College Degree	14.3		15.8	
% More than College Degree	6.8		10.1	
% Urban	66.2		64.9	
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=1,907,701		N=1,853,243	

Notes: Sample is described in notes of Table 1.

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

	Women		Men	
	(1)	(2)	(3)	(4)
Fraction Female, Industry-Occupation	-0.0947 (0.0013)		0.0091 (0.0012)	
Fraction Female, Occupation		-0.0565 (0.0016)		0.0099 (0.0015)
Fraction Female, Industry		-0.0618 (0.0017)		-0.0124 (0.0017)
Mean Male Wage, Industry- Occupation Cell	-0.0012 (0.0001)		-0.0037 (0.0001)	
Mean Female Wage, Industry-Occupation Cell	-0.0003 (0.0002)		-0.0011 (0.0001)	
Mean Male Wage, Occupation		0.0020 (0.0002)		-0.0040 (0.0002)
Mean Female Wage, Occupation		-0.0032 (0.0003)		0.0004 (0.0002)
Mean Male Wage, Industry		-0.0037 (0.0001)		-0.0016 (0.0001)
Mean Female Wage, Industry		0.0030 (0.0002)		-0.0032 (0.0002)
Log Male Wage Variance, Industry- Occupation Cell	-0.0022 (0.0005)		0.0042 (0.0005)	
Log Female Wage Variance, Industry-Occupation Cell	-0.0109 (0.0005)		0.0052 (0.0003)	
Log Male Wage Variance, Occupation		-0.0152 (0.0011)		-0.0047 (0.0010)
Log Female Wage Variance, Occupation		-0.0172 (0.0012)		0.0038 (0.0007)

Log Male Wage Variance, Industry		0.0325 (0.0011)		0.0302 (0.0012)
Female Wage Variance, Industry		-0.0434 (0.0015)		0.0064 (0.0012)
Fraction Female, PUMA	-1.5976 (0.3244)	-1.5684 (0.3254)	-1.1397 (0.3449)	-1.1779 (0.3453)
(Fraction Female, PUMA)^2	1.8177 (0.3308)	1.7896 (0.3307)	0.8490 (0.3471)	0.8840 (0.3475)
Fraction Men Employed, PUMA	-0.6644 (0.0189)	-0.6637 (0.0189)	-0.6567 (0.0168)	-0.6607 (0.0168)
Fraction Women Employed, PUMA	0.3281 (0.0111)	0.3298 (0.0111)	0.3970 (0.0097)	0.3949 (0.0097)
Mean Male Wage, PUMA	-0.0083 (0.0004)	-0.0083 (0.0004)	-0.0034 (0.0003)	-0.0033 (0.0003)
Mean Female Wage, PUMA	0.0084 (0.0006)	0.0086 (0.0006)	0.0052 (0.0006)	0.0055 (0.0006)
Log Male Wage Variance, PUMA	0.0219 (0.0013)	0.0217 (0.0013)	0.0094 (0.0012)	0.0100 (0.0012)
Female Wage Variance, PUMA	-0.0036 (0.0011)	-0.0039 (0.0011)	-0.0002 (0.0010)	-0.0002 (0.0010)
	N=1,907,701		N=1,853,243	

Notes: Sample is described in notes of Table 1. Table reports the results from OLS regressions. Dependent variable is a binary indicator for divorce. Sex-mix variables are the fraction of workers ages 18-55 that are female by occupation, industry and industry-occupation cell. All regressions also include state fixed-effects, state-urban 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: 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	-0.0611 (0.0038)			0.0138 (0.0039)		
Fraction Female, Industry-Occupation (Calculated at state-level)		-0.0354 (0.0037)			0.0295 (0.0036)	
Fraction Female, Occupation (Calculated at state-level)			0.0067 (0.0091)			0.0661 (0.0084
Fraction Female, Industry (Calculated at state-level)			-0.0168 (0.0112)			0.0575 (0.0097
N	1,526,199	1,298,077	1,298,077	1,482,626	1,203,864	1,203,864

Notes: Columns 1 and 4 report results from the regression models used in Columns 1 and 3 of Table 4, with the addition of industry and occupation fixed-effects. Columns 2 and 5 report results from the same industry and occupation fixed-effects model, but with fraction female in industry-occupation, as well as all industry-occupation wage controls, calculated at the state, rather than national, level. Columns 3 and 6 report results based on the model used in columns 2 and 4 of Table 4, with the addition of industry and occupation fixed-effects and with the industry and occupation-specific sex-mix measures and wage controls calculated at the state level. Due to computer memory constraints, samples used in analysis are 80-90% random samples of the eligible observations (details in footnote 13).

Table 6: IV Estimates of Probability of Divorce, 1990 Census

	Women		Men	
	(1)	**(2)**	**(3)**	**(4)**
Fraction Female, Industry-Occupation	-0.4019 (0.0214)		0.2350 (0.0231)	
Fraction Female, Occupation		-1.026 (0.0460)		0.3673 (0.0403)
Fraction Female, Industry		0.6078 (0.0434)		-0.0837 (0.0358)
	N=1,907,701		N=1,853,243	

Notes: Sample is described in the notes of Table 1. Table reports the results from IV regressions. The occupational and industrial composition of PUMA-level employment is used to instrument the fraction female in the occupation, industry and industry-occupation cell as described in the text. All regressions include the wages controls, PUMA-specific controls, state fixed-effects, state-urban effects and individual-specific controls control variables used in the OLS regressions reported in Table 4.

**Table 7: Coefficient on Fraction Female in Industry-Occupation Cell,
Age and Race-Specific Samples**

	Women			Men		
	(1) OLS	(2) FE	(3) IV	(4) OLS	(5) FE	(6) IV
Full Sample (from Tables 4, 5 and 6)	-0.0947 (0.0013)	-0.0611 (0.0038)	-0.4019 (0.0214)	0.0091 (0.0012)	0.0138 (0.0039)	0.2350 (0.0258)
N		1,907,701			1,853,243	
Ages 18-29	-0.0750 (0.0024)	-0.0419 (0.0073)	-0.1599 (0.0595)	0.0087 (0.0028)	-0.0034 (0.0081)	0.2432 (0.0653)
N		415,499			293,932	
Ages 30-40	-0.1062 (0.0020)	-0.0666 (0.0058)	-0.2113 (0.0532)	0.0156 (0.0019)	0.0281 (0.0056)	0.0285 (0.0432)
N		737,606			723,909	
Ages 41-55	-0.0985 (0.0022)	-0.0653 (0.0061)	-0.3312 (0.0718)	0.0151 (0.0018)	0.0156 (0.0054)	0.2103 (0.0363)
N		678,184			737,904	
White	-0.1002 (0.0013)	-0.0615 (0.0038)	-0.4549 (0.0224)	0.0128 (0.0012)	0.0162 (0.0039)	0.2229 (0.0281)
N		1,627,904			1,593,371	
Non-White	-0.0350 (0.0039)	-0.0270 (0.0109)	-0.0007 (0.0413)	-0.0017 (0.0035)	0.0056 (0.0101)	0.1303 (0.0385)
N		259,075			226,337	

Notes: Regression models are the same as those used in columns 1 and 3 of Table 4, columns 1 and 4 of Table 5, and columns 1 and 3 of Table 6 with samples restricted to the described age or race group. 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 8: Descriptive Statistics, NLSY

	Women		Men	
	Mean	St Dev	Mean	St Dev
% Divorced from First Marriage	32.7		28.8	
Age of First Marriage	23.4	(4.6)	24.7	(4.5)
Year of First Marriage	84.3	(5.0)	85.6	(5.0)
Duration of First Marriage in Years (if Divorced)	7.1	(4.5)	6.7	(4.2)
Fraction female in Industry-Occupation at time of Marriage	0.73	(0.26)	0.28	(0.26)
Fraction female in Spouse's Occupation at time of Marriage	0.28	(0.25)	0.71	(0.25)
% Black	17.1		17.9	
% Hispanic	16.2		14.7	
Highest Grade Completed	13.5	(2.5)	13.2	(2.8)
Spouse's Highest Grade Completed	13.4	(2.6)	13.4	(2.3)
% Living with Both Biological Parents in 1979	73.9		73.3	
Number of Children Expected (measured 1979)	2.4	(1.4)	2.4	(1.3)
	N=2,741		N=2,368	

Notes: Unweighted means. Sample is ever-married respondents in the NLSY79 who report an industry and occupation and a spouse's occupation in or before the year of marriage.

Table 9: OLS Estimates of Probability of Divorce, NLSY

	Women		Men	
	(1)	(2)	(3)	(4)
Fraction Female, Industry-Occupation at time of Marriage	-0.1751*** (0.0538)	-0.2433** (0.0832)	-0.0335 (0.0639)	0.0486 (.0946)
Fraction Female, Spouse's Occupation at time of Marriage	0.1536* (0.0680)	0.1149 (0.1218)	-0.0732 (0.0594)	-0.0697 (0.1001)
Industry, Occupation and Spouse Occupation Code Fixed-Effects	1-Digit	2-Digit	1-Digit	2-Digit
N	2,351		2,053	

Notes: Sample is ever-married respondents in the NLSY79 who report an industry and occupation and a spouse's occupation in or before year of marriage. Table reports the results from OLS regressions. Dependent variable is a binary indicator for divorce. Sex-mix variables are the fraction of workers ages 18-55 that are female by occupation and industry-occupation cell calculated from 1980 and 1990 Censuses. All regressions also include industry-occupation and spouse's occupation wage controls, local area sex-mix and economic controls, age of marriage, race indicators (black, Hispanic), highest grade completed, highest grade completed of spouse, indicator for respondent lived with both biological parents at age 14, expected age of marriage in 1979, expected number of children in 1979, state fixed-effects, and state-urban fixed-effects. Initial NLSY 1979 weights are used. *p-value<.05 ** p-value<.01 ***p-value<.001

Table 10: Divorce Hazard Results, NLSY

	Women	Men
Occupation and Industry within past 5 years:		
Fraction Female, Industry-Occupation	-0.7943*** (0.2128)	-0.1668 (0.2693)
Fraction Female, Spouse's Occupation	0.3448 (0.2884)	-0.0417 (0.2223)
Occupation and Industry for Current Year's Work		
Fraction Female, Industry-Occupation	-0.7472*** (0.2168)	-0.2931 (0.2776)
Fraction Female, Spouse's Occupation	0.0520 (0.2786)	-0.1041 (0.2511)
Worked this Year	1.039 (0.6280)	-1.181 (1.014)
Spouse Worked this Year	-0.3050 (0.3097)	0.1449 (0.3821)
N	33,436	26,711

Notes: Sample is ever-married respondents in the NLSY79. Table reports estimates from a logistic model for divorce. Sex-mix variables are the fraction of workers ages 18-55 that are female by occupation and industry-occupation cell calculated from 1980 and 1990 Census. Sex-mix measures are set to zero if the respondent does not work in a given year. All regressions also include industry-occupation and spouse's occupation wage controls, local area sex-mix and economic controls, age of marriage, race indicators (black, Hispanic), highest grade completed, highest grade completed of spouse, indicator for respondent lived with both biological parents at age 14, expected age of marriage in 1979, expected number of children in 1979, state fixed-effects, state-urban fixed-effects, year fixed-effects, duration of marriage indicators, 1-digit industry code fixed-effects, 1-digit occupation code fixed-effects and 1-digit spouse occupation code fixed-effects. Initial NLSY 1979 weights are used. *p-value<.05 ** p-value<.01 ***p-value<.001

Appendix

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.