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

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Sexually-Integrated Workplaces and Divorce:

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 consid71 tdppi8eaw3.9195 4om 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 tim 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]

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[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.

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 We299912 17po -0.0008op463.m(7.7599 Tm(i)Tj12 0 -0.008op463.m(7.7599 Tm(i)Tanh4j12

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

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:

$$\begin{aligned} Y_{ionls} \quad {}_{0} \quad {}_{1}FrFemOCC_{o} \quad {}_{2}FrFemIND_{n} \quad {}_{3}FrFemLOC_{l} \quad {}_{4}(FrFemLOC_{l})^{2} \\ {}_{5}FrMEMP_{l} \quad {}_{6}FrWEMP_{l} \quad X_{ionls} \; {}_{7} \quad STATE_{s} \quad (STATE_{s} * Urban_{i}) \quad {}_{ionls} \end{aligned} \tag{1}$$

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 22 0)02 0 0 10.02 127.94ide877.770.02 0 0 108he s 70.02 0 0 108g ihoeffd ev0.0. 3929.20.3king a2 0 0 10.02 110.9304

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.

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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 ma 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 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 mo

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:

$$IVOCCMale_l \qquad {}_{o} \frac{ShareMaleEmp_{ol} * FrFemOCC_o}{TotalMaleEmp_l}, \tag{2}$$

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 $TotalMaleEmp_l \qquad {}_{o} ShareMaleEmp_{ol}$ is total male employment in PUMA l. An analogous instrument can be calculated for the fraction female in a male worker's industry:

$$IVINDMale_l \qquad {}_{n} \frac{ShareMaleEmp_{nl} * FrFemIND_n}{TotalMaleEmp_l}, \tag{3}$$

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 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.

[19] Glass, Sherry. 2003.

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Table 1: Distribution of Fraction

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 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)

2Log Female 0Wa T j E T E N

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 (0.0016T/TT1 1 Tf0 Tc 12 0 0 12 tF62 3.b64Tc 016T/TT1 1 T		

Table 6: Fixed-Effects Estimates of Probabilit

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 Ind[redacted]pation Cell,
Age and Race-Specific Sam[redacted]

	Women		[redacted]**en**	
	(1) OLS	**(2) IV**	[redacted]	**(4) IV**
Full Sample (from Tables 5 and 7)	-0.0948 (0.0012)	-0.4028 (0.0207)	[redacted]	0.2309 (0.0456)
N	1,907,701		[redacted],243	
[redacted]				
Ages 30-40	-0.1062 (0.0020)	-0.2113 (0.0532)	[redacted]	0.0285 (0.0432)
N	737,606		723,909	
Ages 41-55				

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	