DO WAGES FALL WHEN WOMEN ENTER AN OCCUPATION?

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ABSTRACT

I present the first causal evidence on the effect of the entry of women into occupations on the wages of those occupations. To determine the causal effect of a change in gender composition, I construct a shift-share instrument that interacts the dramatic increase in the relative educational attainment and workforce participation of women from 1960-2010 with the relative likelihood of men and women to enter the occupation. I find that a 10 percentage-point increase in the fraction of females within an occupation leads to an 8 percent decrease in average male wage and a 7 percent decrease in average female wage in the concurrent census year. Over the 10 years following the change in the gender composition, I find that the effect of such an increase in the fraction of females grows to a 9 percent decrease in male wages and an 13 percent decrease in female wages. I present suggestive evidence attributing this finding to effects of gender composition on the prestige and amenity value of occupations.

JEL Codes: J16, J24, J31, J71, J82, J110

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Section I: Introduction

W~~Among full-time workers, w~~omen earn 19% less than men on average among full-time workers (Blau & Kahn, 2016). While this ~~wage gap has closed~~difference has decreased since the 1960s, it has changed less than would be expected based on women's relative gains in education and experience. This persistent wage gap can be accounted for primarily by the fact that women work in lower-paying occupations and industries than do men, with occupation and industry accounting for 51% of the wage gap in 2010 (Blau & Kahn, 2016). As a result, understanding why women work in lower-paying occupations than men is crucial to understanding why women earn less than men. This paper presents the first empirical evidence that the presence of women in an occupation results in declines in average wages for both men and women in that occupation. This finding suggests that differences in pay between male and female dominated occupations may reflect ways in which the gender of workers changes the nature and compensation of work.

There are two broad explanations for the negative relationship between women's representation and pay in occupations. One explanation is that differences in the occupations of men and women reflect differences in the constraints, preferences and abilities of male and female workers. In this case, women's greater household responsibilities would lead female workers to choose less demanding, lower productivity, and lower paying jobs relative to men with equal levels of human capital (Becker, 1985).[1] Alternately, discrimination may push women to work in lower-paying occupations than do similar men, because women may face greater barriers in high-paying male-dominated occupations than they do in similarly skillful but lower-paying female-dominated occupations. These barriers can take many forms, such as lower likelihood of promotion (Thomas, 2019; Maume, 1999), reduced access to mentorship and networking (Chen, et al., 2015), or higher rates of discrimination and sexual harassment (Parker, 2018).

[1] Relative to men, women are more likely to choose occupations that allow career interruptions (Görlich & De Grip, 2007), do not require long or inflexible working hours (Goldin C. , 2014), and do not require intense competition or uncertain rewards (Niederle & Vesterlund, 2005), leading to lower wages.

Both of these explanations examine why otherwise similar male and female workers may choose differently among the bundles of wages and amenities offered by different occupations. However, if the presence of women in an occupation causes a decline in that occupation's wages, any difference in the way that men and women choose occupations will lower average female wages relative to male wages. In turn, a failure to recognize this effect could lead policymakers to overestimate the benefits of attempting to change the occupation decisions of women.

While a large sociology literature shows a negative correlation between changes in average male and female wage and changes in the female share of an occupation's workforce[2], this work is exposed to significant potential biases. Most significantly, changes in the gender composition of an occupation may be related to changes in returns to skill in that occupation, to changes in the returns to experience or long hours, or to changes in the relative size of male-dominated and female-dominated specialties or sub-occupations.

I present evidence on the effect of changes in the gender composition of an occupation on ~~the average wage of an occupation~~wages that more plausibly handles the selection problems that have plagued earlier research. I estimate a panel regression of the log average male and female wage of older workers in an occupation on the fraction of female workers among younger workers in that occupation[3]. Because changes in the gender composition of an occupation may be driven by factors that influence current and future wages[4], I construct a shift-share instrument for gender composition identified from the dramatic increase in relative female labor force participation and educational attainment from 1960 to 2010,

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[2] Previous studies have either depended on cross-sectional variation in gender composition and wage (Cohen & Huffman, 2003), (England P. , 1992), on time-series variation (Catanzarite, 2003), (England, Allison, & Wu, 2007), (Karlin, England, & Richardson, 2002), or on panel data (Levanon & Allison, 2009). (Tam, 1997) shows that these methods are sensitive to changes in the set of included covariates.

[3] I separate older and younger workers in order to account for the possibility that the characteristics of the male and female labor force in each education group change as the growth of women's work and education shifts the margin of work and college attendance.

[4] Changes in the fraction of workers who are female in an occupation can be related to both endogenous increases and decreases in wage. If occupations that become less demanding see ~~increases~~ decreases in wages and increases in female labor supply, ~~changes in occupations~~this change will introduce negative bias to panel regressions. On the other hand, if women's increased education and workforce engagement leads to increases in the fraction female in occupations with increasing skill requirements, changes in occupations will introduce positive bias to panel regressions.

measured using the decennial census. Over this time, the female share of educated workers under the age of 35 increased rapidly, particularly for advanced degree holders. This change is unlikely to be attributable to changes in the wages of particular occupations—previous research has attributed this rise in women's work and education to increased access to contraceptives, changing cultural attitudes towards female labor force participation, and changes in the administration of high schools and universities (Goldin, Katz, & Kuziemko, 2006).

This increase in female educational attainment increased the availability of women in all occupations that primarily hired workers with college and advanced degrees. However, because men and women of similar educational attainment ~~tended~~ tend to work in different occupations, some occupations were more exposed to this change than others. ~~In particular, the rise of female educational attainment and labor force participation induced larger changes in the female fraction of workers in o~~Occupations that were equally popular with educated men and women in 1980, such as psychology, had larger induced changes in the female fraction of workers than did ~~than in~~ occupations that were disproportionately popular with either men or women in 1980, such as physics or nursing. ~~The increased availability of highly educated women led to the largest changes in gender composition in o~~Occupations chosen equally by men and women had the largest induced change in gender composition because female-dominated occupations had little scope to increase their fraction female, while few of the newly-educated female workers sorted into male-dominated occupations.[5]

I use this difference across occupations in exposure to the national trend of increased relative educational attainment of female workers to construct a shift-share instrument for the gender composition of each occupation (Bartik, 1991). This instrument estimates the fraction of available workers in each

[5] For example, consider an occupation that employs only workers with post-graduate degrees. If men and women with post-graduate degrees were equally likely to work in the occupation, it would have seen an increase in percent female from 15% in 1960 to 59% in 2010, matching the rise in the female share of post-graduate degrees. On the other hand, if post-graduate women were 20 times as likely to work in the occupation as were post-graduate men, the occupation would have been nearly 80% female even in 1960, and nearly 97% female in 2010. Likewise, if post graduate women were 1/20 as likely to work in the occupation as were post-graduate men, the occupation would have been less than 1% female in 1960, but still only 7% female in 2010.

occupation who would be female if the proportion of workers with each education level, as well as the likelihood that a male or female worker of each education level worked in the occupation, remained fixed to their 1980 levels, while the gender composition of each education level varied over time. Because it is not identified off of changes in the occupational preferences of men and women, this instrument accounts for the endogenous relationship between changes in occupational preferences and changes in wages.

I find strong evidence that increased female representation in an occupation leads to lower average wages for men and women, both in the immediate term and in the longer term. I estimate that a 10 percentage-point increase in the female share of an occupation's workforce leads to an 8.2% decline in average male wage, and a 7.2% decline in average female wage, measured contemporaneously. Over the following ten years, the effect grows to an 8.7% decline in average wage for males and a 13.0% decline in average wage for females. Over the following 20 years, the effect on wage falls to a 4.5% decline in average male wage and a 8.6% decline in average female wage. These results are large compared to the cross-sectional relationship between gender composition and pay. In 2010, when controlling for age and education, a 10 percentage point increase in female share was associated with a 5% decrease in average wage for men and a 4% decrease in average wage for women (Supplementary Table A.I).

While this approach faces far fewer threats to identification than do panel approaches that do not instrument for fraction female, several potential sources of endogeneity must be accounted for. First, the increase in women's work and education that led to disparate changes in the gender composition First, the increase in women's work and education that led to disparate changes in the gender composition of occupations also had disparate effects on labor supply to those occupations. As a result, I must disentangle the labor supply effect of women's work and education decisions from the gender composition effect of those decisions. I do this by constructing a control for "induced labor supply"—that is, for changes in labor supply that would have resulted from changes in men and women's work and education decisions, were men and women's occupation choices unchanged. This control accounts for the fact that the rise in women's education and labor force participation results in

~~decisions, holding occupation choices constant.~~ This control accounts for the fact that the rise in women's education and labor force participation results in higher labor supply in occupations ~~in which a high fraction of workers are highly educated women~~that predominantly employ highly educated women.[6] Because the occupations with the greatest predicted changes in gender composition are gender-balanced, while the occupations with the greatest predicted change in labor supply are female-dominated, this labor supply control is not collinear with the gender composition instrument.

Second, while the use of this instrument addresses many of the confounding relationships between changes in an occupation's gender composition and changes in wage, shift-share instruments may be confounded by characteristics of occupations that are correlated with the base-year occupation decisions of men and women. Of particular concern is the possibility that the relative likelihood of female employment in an occupation in 1980 is related to the occupation's skill requirements. In this case, changes in the returns to skills over the last half-century could affect the wages of occupations with large predicted changes in gender composition differently than they affect the wages of occupations with small predicted changes. Work by Autor, Levy and Murnane (2003) and Deming (2017) have identified a few key dimensions of skill that have seen changing returns over this period, with the returns to social skills, math/analytical skills and service increasing and the returns to routine tasks decreasing. While these skills are weakly related to the gender composition of occupations, with social skill requirements highest and math/analytical skills lowest in female-dominated occupations, there is substantial variation in the skill requirements of occupations with similar gender ratios. The negative effect of increased percent female on wage is robust to the inclusion of time-varying skill effects, indicating that changing returns to skills identified as important in previous literature do not explain the estimated relationship.

[6] Increases in the fraction of the workforce consisting of highly educated women would be expected to increase the labor supply of female-dominated occupations so long as demand for educated women did not grow as quickly as supply for educated female workers. This hypothesis is supported by the presence of several supply-side factors that explain the rise in women's work and education, such as delays in childbearing, increased work expectations among women, and greater high-school preparedness (Goldin, Katz, & Kuziemko, 2006). Consistent with a supply-side explanation for the rise in women's work and education, the labor supply index is positively correlated to employment and negatively correlated with wages. If instead changes in demand for female workers drove women's work and education, the labor supply index would be negatively correlated employment and positively correlated with wages.

I make several contributions to the literature. First, I contribute to a related literature on the effect of working in a female-dominated occupation on the wage of an individual worker, pioneered by Hirsch and Macpherson (1995) and updated by Addison, Ozturk and Wang (2018).[7] Hirsch and Macpherson measure the effect of the gender composition of a worker's occupation on wages for workers who changed occupations in the CPS. While this analysis finds evidence that workers received lower pay in occupations with a higher share of female workers, the difference is mostly explained by measured job characteristics, including training requirements and occupational hazards. This work provides compelling evidence that workers on the margin between male-dominated and female-dominated occupations do not require a compensating differential for working in the male-dominated occupation. However, this research design cannot speak to the question of whether the process by which wages are determined differs between male and female dominated occupations. For instance, if returns to some skills are lower in female-dominated occupations, this will affect which workers are on the margin between these occupations and male-dominated occupations, but will not affect the wage penalty for working in the female-dominated occupation for marginal workers. ~~Because an analysis of job switchers necessarily focuses on workers on the margin between two occupations, this finding suggests that women do not require a compensating differential for working in male-dominated occupations.~~

I expand on Hirsch and Macpherson's analysis by examining the effect of gender composition on wage at the occupation level, rather than at the individual level. If the wage level falls for all workers in an occupation, workers on the margin with other occupations will exit, while non-marginal workers will experience a wage decline. Because a new set of workers will be on the margin with other occupations after this wage decline, this decline in overall wage rates will not appear as a wage penalty in the Hirsch and MacPherson strategy. It will, however, be captured by an analysis of average wages.

[7] Other important work in this literature includes Murphy and Oesch (2015), who examine the wage penalty of changing to a female-dominated occupation in the UK, Germany and Switzerland, and Pitts (2002), who structurally estimates the wage offers of workers in female-dominated occupations, were they to work in male-dominated occupations,

Next, I contribute to a growing literature examining the effect of changes in gender composition on the characteristics of occupations. Previous research has hypothesized that the entry of women into occupations causes declines in an occupation's prestige (Goldin C. , 2014), a hypothesis supported by the exit of men from occupations that pass a threshold percentage female (Pan, 2010). In addition, Goldin and Katz (2011, 2012) find that the gender composition of an occupation coevolves with returns to long hours and rates of independent firm ownership, and they speculate that some of that coevolution may be due to the effect of a growing female workforce on the organization of occupations. The large, negative effect of gender composition on wage found in this paper suggests that these hypothesized effects of gender composition on occupations may be large and have significant wage consequences.

The rest of this paper is organized as follows: Section 2 introduces theory suggesting that gender composition of an occupation can affect its wage. Section 3 describes the data used in this analysis. Section 4 describes the empirical strategy of the paper, Section 5 describes the results, Section 6 describes tests of the mechanisms of effect, and Section 7 concludes.

Section 2: Theory

Gender composition can affect the average wage of an occupation through three primary channels: changes in the characteristics of workers in an occupation, changes in the characteristics of the occupation, and changes in the way that worker characteristics are compensated. That is, a change in the percent of workers who are female can cause workers to enter or exit an occupation (e.g. highly skilled workers might exit the occupation), can cause an occupation to change amenities (e.g., the occupation might require shorter working hours), or can change the returns to worker characteristics (e.g. returns to particular skills could increase or the cost of career interruptions could decrease). In this analysis, all three of these channels contribute to the causal effect of interest.

Those causal channels may in turn be activated by several underlying economic mechanisms. I will divide these mechanisms into two broad categories: the effect of prestige/perceptions of the occupation on wages, and effects of amenities/characteristics of the occupation on wages.

Mechanism 1: Prestige/Perceptions:

Declines in wage as a result of increased fraction female may be the result of negative perceptions of women's abilities or work engagement influencing the perception of women's occupations. For example, Goldin (2014) shows that if observers outside an occupation believe that female workers have lower abilities on average than male workers, and they cannot perceive the true skill requirements of an occupation, male dominated occupations will be perceived to require more skill than do female-dominated occupations. As a result, an increase in female representation in an occupation may result in a decline in perceived difficulty (prestige) of the occupation among individuals not working in the occupation.

This can affect wages of workers both through direct channels, like declines in demand for labor in the occupation, and through changes in the composition of the workforce. Workforce composition may change if, in addition to its psychic benefits, prestige influences the ability of a worker to get a higher-paying job in the future. Because workers change occupation frequently, with between 20-30% of respondents in the NLSY reporting a change in occupation and firm each year (Sullivan, 2010), the signal of ability sent by current occupation may be an important factor in determining future wages. If lost opportunities for career advancement are most costly to high ability workers, declining prestige in an occupation due to an increasing share of female workers will result in selection of lower-ability male and female workers into that occupation. As a result, even if compensating differentials increase wages in the occupation conditional on ability, average wage may fall due to a shift in the composition of the workforce toward lower-ability or lower-ambition workers.

The effect of a decline in prestige on the wages in an occupation may be long-lasting. Basu (2017) notes that if many agents in society benefit from giving preferential treatment to groups that are preferred by others, discrimination can persist in equilibrium. For example, suppose that technology companies can be run either by engineers or by marketers. An increase in the percent of marketers who are female leads investors and potential employees to conclude that marketing is a less demanding and less prestigious occupation than is engineering. If, as a result, employees prefer to work for companies run by engineers and investors prefer to invest in companies run by engineers, engineers will be favored for leadership positions over marketers. Once this equilibrium is established, employees will prefer to work for engineers, because engineers are better at securing funding than are marketers. Investors will prefer to invest in engineer-run companies, because engineers can secure better employees. Equilibrium outcomes of this sort could mean that occupations that see an increase in the share of female workers become unable to attract ambitious workers in the long run.

Mechanism 2: Changing amenities for workers in the occupation:

Increased female representation in an occupation can also change wages by changing the provision of costly job amenities if women value the amenities differently than men. In particular, women on average pay higher costs for long and inflexible working hours than do men. In particular, women on average pay higher costs for long and inflexible working hours than do men (Goldin C. , 2014), (Mas & Pallais, 2017), are less willing to apply for jobs with competitive compensation schemes (Flory, Leibbrandt, & List, 2015), and are less willing to accept jobs that risk fatality (DeLeire & Levy, 2004).

In many circumstances, work can be organized in a way that allows shorter and more flexible hours, reduced competition, less accident risk or greater benefits at the cost of lower productivity or higher costs. In these circumstances, an increase in the availability of female workers may cause firms to reorganize in a way that reflects the preferences of female workers. Lee and Thompson (2019) construct and empirically test a model with this implication, where firms offer workers individual wages and a benefits package that is common to all workers. If women value benefits more highly than do men on average, they show that firms will provide more benefits and lower wages when their industry has a larger share of female workers.

Goldin and Katz (2011) provide a specific example of this process in the field of veterinary medicine. Small veterinary practices can reduce the costs of flexibility in hours by referring emergencies to regional veterinary hospitals, rather than by keeping staff “on call” outside of normal working hours. A veterinary practice will choose to refer emergencies if the cost of lost productivity is lower than the benefit of lower wages, due to the amenity value of flexible hours. Because female veterinarians are less likely than male veterinarians to work overtime hours (Goldin & Katz, 2011), referring cases becomes more profitable as the percent of veterinarians who are female increases. In many cases, such reorganizations reflect the gender composition of occupations as much as they do particular firms—as the fraction of veterinarians who are female increases, regional hospitals will receive more referrals, become more prevalent, and thus suffer fewer disadvantages relative to on-call hours. Indeed, as women began to dominate the veterinary profession—making up 80% of recent veterinary school graduates—emergency services have shifted to regional hospitals (Goldin & Katz, 2011).

This process can occur for any amenity that is valued more by female than by male workers on average. Increased amenities will lower wages conditional on worker characteristics by increasing costs and/or lowering productivity. In addition, it will make the occupation more appealing to workers with a high value of the amenity (i.e. those who face high costs from long hours, competition etc.) and less appealing to workers with a low value of the amenity, due to decreased wages. If workers who face greater costs from competitive pay structures, long hours and other amenities are on average less productive or less ambitious than those who face low costs, lower productivity workers will sort into an occupation as the workplace amenities improve.

It is important to note that these two mechanisms are not mutually exclusive, and in fact may be mutually reinforcing. A decline in an occupation's prestige may result in a change in the composition of a workforce toward less ambitious workers, who in turn may demand flexible working hours, less risky compensation schemes and similar amenities. Likewise, higher provision of amenities in an occupation may result in lowered perceptions of the occupation's difficulty, and thus diminished prestige.

Section III: Data

The main analysis in this paper is conducted using the IPUMS microdata files of the Decennial Census (1960-2000) and the American Community Survey (2009-2014) (Ruggles, Genadek, Goeken, Grover, & Sobek, 2015). Supplementary analysis is performed using the Panel Study of Income Dynamics (1968-2016) and the Current Population survey (1968-2016). These data are described in detail in the Data Appendix, but I focus on a few important issues here.

First, I choose the years 1960 to 2010 in order to maximize sample size while maintaining consistent ten-year intervals between included years. I exclude the 1950 census because the sampling technique used to measure occupation was non-random within households selected for the long-form census. In particular, the 1950 census asked about occupation only for one member of each household, selected by the household. While it is possible to reweight this sample in order to make it nationally representative, doing so cannot account for correlations between the likelihood that the selected respondent is female and the occupation of the female household member.

Second, I define occupation using a modification of the 1990 occupation codes constructed by IPUMS. These codes are constructed using crosswalks to earlier and later occupation coding schemes, as described by Myer and Osborne (2005). I deviate from the 1990 harmonized codes in the 2000 and 2010 census by using the gender-specific occupation classifications provided by Blau, Brummund and Liu (2012). Blau, Brummond and Liu (2012) observed that because the most prevalent 1990 codes for each 2000 code differ for male and female workers, gender segregation of occupations appears to decrease in 2000 when measured with harmonized 1990 occupation codes. To address this, Blau, Brummond and Liu (2012) define a gender-specific crosswalk between the 1990 and 2000 codes by choosing the most prevalent 1990 occupation separately for women and men for each 2000 occupation code.

Finally, I include all occupations in my analysis, regardless of size, and do not weight occupations by size. There are two reasons that one might weight occupations in this context. First, as discussed in Solon, Haider and Woolridge (2015), the measurement error associated with mean wage in each occupation is decreasing in the size of the occupation, causing heteroskedasticity in the error term. Were errors uncorrelated within group, weighting occupations by the square root of their sample size would address this heteroskedasticity. However, Solon, Haider and Woolridge (2015) note that when errors are clustered within occupation, as they are in this analysis, unweighted estimates may be more precise. I performed a simulated power analyses, considering four ways of addressing unequal size of occupations: no weights or exclusions, weighting by sample size, weighting by the square root of sample size, and eliminating occupations represented by fewer than 100 workers. Of these, analyses with no weights or exclusions had the most power. Secondly, weighting estimates by the number of workers in each occupation would make the estimated effect representative of the effect in the average worker's occupation, whereas unweighted estimates represent the effect in the average occupation. Because occupations, rather than workers, are the subject of this analysis, I perform an analysis that is representative of the average occupation. As a robustness check, I present estimates weighted by the number of workers in each occupation Supplementary Table A.II. when weighting by number of workers, the estimated effect of fraction female on wage is smaller than when not weighting, but still large and statistically significant.

Section IV: Empirical Strategy

In order to motivate the approach taken in this paper, I consider a hypothetical experiment testing the effect of gender composition on wage. One such experiment might divide occupations into treatment and control, and would retire a random subset of the male workers in the treated occupations and replace them by cloning a random subset of female workers. The experimenter would then measure average wages for male and female workers in this occupation, including in the average workers who enter the occupation after the cloning experiment and excluding in the average workers who exit the occupation after the cloning experiment.

This experiment has a few important properties. First, it would effectively replace male workers whose characteristics are distributed typically for males working in the occupation with female workers whose characteristics are distributed typically for females working in the occupation. As a result, an effect on wage in this experiment could come through direct effects of gender, such as declines in prestige for occupations with a large female workforce, but could also come through indirect effects of gender-linked characteristics such as demand for amenities. Second, because the experimental effects include entry to and exit from the occupation, the estimated coefficient is an effect on the average wage of the occupation, not on the expected wage of a worker with a given set of characteristics.

It is important to note that this experiment measures only one of several parameters that could be considered an effect of the gender composition of an occupation on wage. Because this experiment changes gender composition without changing the preferences for or barriers to an occupation, it may have quite different effects than would an experiment that provided a quota for female workers in an occupation, that lowered discrimination against women in an occupation, or that implemented affirmative action in education that is a prerequisite for work in the occupation. In particular, these alternative experiments are likely to include labor supply effects through the overcrowding channel (Bergmann, 1974) that are not included in the effects measured in the primary experiment. Yet another alternative might be to transform the sex of some subset of workers in treated occupations from male to female, without changing the preferences, beliefs or constraints of the workers at all. Doing this would present an effect of gender composition absent the effect of any differences between male and female workers within an occupation.

4.1: Baseline Panel Regression:

My goal is to replicate the primary experiment described above by examining the effect of changes in the fraction of workers in an occupation who are female in one census year on wages in future census years

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for male and female workers, indexed by gender *g*, and for zero, ten and twenty-year time-lags, indexed by *k*:

$$W_{g,j,t+k} = \alpha + \beta_{1k} f_{j,t} + \beta_{2kf} W_{f,j,t} + \beta_{2km} W_{m,j,t} + \delta_{fk} X_{f,j,t} + \delta_{mk} X_{m,j,t} + \gamma_{jk} + \sigma_{tk} + \varepsilon_{g,j,t+k} \quad (1)$$

Where $W_{g,j,t}$ is the log mean wage of workers aged 45-65 with gender g in occupation j at year t, $f_{j,t}$ is the proportion female of workers aged 22-65 in occupation j, and $X_{m,j,t}$, $X_{f,j,t}$ are the mean workforce characteristics male and female incumbent workers respectively. In this regression, β_{1k} measures the effect of fraction female in year t on log mean wage in year $t+k$, and includes the effects of changing workforce composition (other than characteristics in $X_{f,j,t}$), changing returns to worker characteristics and changing occupation characteristics. I highlight a few key characteristics of this regression below.

Because the decision to work and attend college is changing for men and women over this period, the unobserved abilities of workers of each education level may have changed (Lovenheim & Reynolds, 2011). I address this by estimating the effect on wages of workers aged 45-65, who I label "Incumbent" workers, from changes in the gender composition of the occupation driven by the education and work decisions of workers aged 22-35, who I label "Newcomer" workers.

Because β_{1k} includes the effect of fraction female on workforce composition and occupation characteristics, controls for time-varying occupation characteristics and worker characteristics may attenuate estimates by capturing changes in those characteristics induced by the change in gender composition. As a result, I include in the controls $X_{j,t}$ only a limited set controls for the age and the distribution of education for workers in the occupation, because these are the occupation characteristics most likely to be confounded with the instrument. In particular, I control for the percent of workers between the ages of 45-55, the percent of male and female workers aged 55-65, and the average age of the occupation among workers aged 45-65, and the average age squared among workers aged 45-65. I also control for the percent of male and female workers who have less than a high school education, a high school education, an associate's degree, a bachelor's degree, and an advanced degree.

Next, I estimate a linear relationship between fraction female and log wage. I do this for a mix of practical and theoretical reasons. A linear relationship is supported by theory because several potential mechanisms suggest that wages should be influenced by changes to the percent of female workers, rather than to the growth rate of female employment. In particular, the returns to amenities that operate as public goods in an occupation depend on the sum of utilities across all workers, so differences in the value of those goods to men and women should be weighted by their raw percentages. The proper functional form of the relationship between percent female, prestige and wage is less clear, but absolute changes in percent female are observable to those outside an occupation in a way that proportional changes may not be.

As a practical matter, examining changes in the percent female avoids measurement error. The decennial census is a very large dataset, sampling 1% of the US population in year 1960, 1970, 2000 and 2010, and sampling 5% of the US population in 1980 and 1990. However, construction of the shift-share instrument (described in detail in section 4.2) is data intensive, and places a great deal of weight on single individuals, particularly in heavily male-dominated and heavily female-dominated occupations. This is particularly true for occupations that were small in 1980 (the base year). An occupation with 200 members that is 2% female would be represented in the base-year by only four female workers. As a result, measurement of the educational composition of female workers in that occupation is necessarily extremely noisy. While the predicted percent changes in such an occupation would be small regardless of noise in the number of females and educational distribution of those females, changes in log percent female are large and highly variable for such an occupation.

Finally, I examine changes in wage over the course of 20 years because occupations are likely to evolve slowly in response to changes in gender composition. Firms may take time to provide amenities to their workers after demand for those amenities increases. Furthermore, the wage consequences of those changes, due both to selection into and out of the occupation and to the productivity losses associated with occupational amenities, may differ in the immediate-term from the long-term. Likewise, negative prestige consequences of increased female representation in an occupation may take time to arise, as those outside of the occupation observe the change in the workforce, and may change in importance over time as observers get additional information about the occupation.

4.1.1 Potential Confounders

Because there are several important sources of endogeneity between fraction female and wage, equation 1 does not identify the causal effect of fraction female on wage. A few particularly important sources of endogeneity are described below.

Female Labor-Force Engagement: Because female attachment to the labor force grew considerably from 1960-2010, and because women dramatically increased their education levels relative to men, occupations that attract new and highly educated workers would be expected to see increases in the representation of female workers, relative to occupations that are less attractive to new and highly-educated workers. Because new, educated workers are likely to choose occupations that are expected to experience high growth and a positive wage outlook, this will produce a positive relationship between changes in percent female and average wage. In addition to the mechanical effect—more new, educated workers leading to more female workers because new, educated workers are disproportionately female—we might expect that discrimination against female workers is less costly in occupations that are not experiencing growing labor demand and tight labor markets. In addition, women's high elasticity of labor supply relative to men (McClelland, Mok, & Pierce, 2014) implies that changes in the wage level of an occupation will have a larger effect on the labor force participation of women trained to work in the occupation than on the labor force participation of men trained to work in the occupation, causing positive correlation between wage and percent female.

Returns to Skills Within an Occupation: Because men and women have different average skills, changes in the returns to skills may also affect both gender composition and wage growth. If the returns to skills that are more prevalent among male workers decrease in an occupation (for instance, if physical strength becomes less important), wages will fall and female representation will rise. Meanwhile, if the returns to skills that are more prevalent among female workers increase in an occupation (for instance, if social skills become more important (Deming, 2017)), wages and female representation will rise. This bias could be both positive or negative, depending on the relative importance of and changes in disproportionately male and disproportionately female skills.

Aggregation of Sub-Occupations: Finally, because occupation definitions necessarily aggregate over several types of work, changes in the relative size of male-dominated and female-dominated sub-occupations can change both gender composition and wages. For example, secretaries at manufacturing firms are more likely to be male and are more highly paid than are secretaries at other firms. A decline in the percent of secretaries who are manufacturing secretaries will thus increase the percent female among secretaries and decrease wages, even if it is not associated with any change in wage or gender composition among manufacturing secretaries or among other secretaries.

4.2: Instrumenting for Gender Composition

4.2.1: Definition of Primary Instrument

I address these confounders by exploiting the increase in the female fraction of the workforce brought about by the rapid increase in women's relative educational attainment and labor force participation from 1960 to 2010. As shown in Figure I, the percent of college degree and advanced degree holders who are female increased dramatically among workers aged 22-35, with just over 11% of advanced degree holders in the workforce female in 1960, compared to nearly 60% in 2010.

[Figure I]

Were all occupations to remain fixed at some point in time in their relative attractiveness to male and female workers with each education level, this change in the educational attainment of female workers relative to male workers would increase the availability of female workers in all occupations, particularly in occupations that hire workers with college degrees or advanced degrees. However, the extent of the implied change would depend on the relative prevalence of the occupation among males and females of each education level. Occupations like physics or nursing, that are predominantly chosen by workers of one gender, would have small induced changes in gender composition, while occupations like psychology, which are chosen by both male and female workers, would have larger induced changes. The occupations for which these changes are largest and smallest over the full sample period are shown in Supplementary Table A.III and A.IV.

The occupations that experienced the largest change in gender composition induced by the instrument are diverse. Social sciences are highly represented, with Psychologists, Economists and Social Scientists (not otherwise classified) all among the ten occupations with the highest predicted change in fraction female. Relatively male-dominated segments of the education field are also well represented, with High School Subject Instructors and Managers in Education in the top 10. Therapy and Social Work also experienced high predicted growth in fraction female, as did writers and authors, editors and reporters and artists/entertainers. With the exception of artists/entertainers, workers in these occupations were highly educated in 1980, with 70-95% of workers holding at least a college degree. They are all people-oriented, involving either close contact with other people or careful study of human behavior. However, they vary substantially in other skills and attributes, including occupations with different firm structures, risk and job requirements.

Occupations that experienced the smallest change in gender composition induced by the instrument are dominated by engineering occupations, but also include particularly male-dominated and particularly female-dominated healthcare occupations. Six of the ten occupations with the smallest predicted change in fraction female are engineering occupations, as are seven of the top 20. All of these engineering occupations had few female workers in 1980 and required high levels of education. Other occupations with a low predicted change in fraction female are quite diverse, however, including registered nurses, dieticians, physicists and dentists. While registered nurses and dieticians were overwhelmingly female in 1980, dentists were overwhelmingly male—only 5% of dentists were female in 1980. The low predicted changes in dentistry and optometry contrast with other health diagnosing occupations—pharmacists, physicians and veterinarians were all more prevalent occupations for women than were dentistry and optometry in 1980, and were all predicted to have average or above-average increases in fraction female.

I define a shift-share instrument (Bartik, 1991) to capture these induced gender compositions by fixing in 1980 the fraction of workers in each occupation who have each of five education types, defined by highest degree attained: less than high school, high school, some college, bachelor's degree, and master's degree or higher. In addition, I fix the fraction of men with each education type working in the occupation and the fraction of women with each education type working in the occupation. The instrument gives the fraction of workers in occupation j who would be female were these base-year fractions fixed, but the gender composition of each education level varied from year to year. I define the instrument as:

$$\widetilde{f_{J,t}} = \sum_A \gamma_{a,1980}^{j} * \left(\frac{\omega_{aj,1980}^{F} * f_{a,t}^{\bar{J}}}{\omega_{aj,1980}^{F} * f_{a,t}^{\bar{J}} + \omega_{aj,1980}^{M} * \left(1 - f_{a,t}^{\bar{J}}\right)} \right) \tag{2}$$

Where 1980 is the base year, $\omega_{aj,1980}^{F}$ is the fraction of females with educational attainment a in occupation j in 1980, $\omega_{aj,1980}^{M}$ is the fraction of males with educational attainment *a* in occupation *j* in 1980, $f_{a,t}^{\bar{J}}$ is the female fraction of workers with educational attainment *a* in year *t*, excluding those employed in the same occupation as occupation j, and $\gamma_{a,1980}^{j}$ is the fraction of workers in occupation *j* with educational attainment *a*. This can be simplified by defining $\Omega_{a,j,1980} = \frac{\omega_{aj,1980}^{F}}{\omega_{aj,1980}^{M}}$, and $F_{a,t}^{\bar{J}} = \frac{f_{a,t}^{\bar{J}}}{1 - f_{a,t}^{\bar{J}}}$. I can then rewrite equation (2) as:

$$\widetilde{f_{J,t}} = \sum_A \gamma_{a,1980}^{j} * \left(\frac{\Omega_{a,j,1980} * F_{a,t}^{\bar{J}}}{\Omega_{a,j,1980} * F_{a,t}^{\bar{J}} + 1} \right) \tag{3}$$

To illustrate the predictions of this instrument for male-dominated, female-dominated and balanced occupations, consider the examples of physics, psychology and nursing. Table I gives the fixed base-year values for each of these occupations. Physics was a far more common occupation for males of each education level than for females of the same education level in 1980, with a man holding an advanced degree more than 10 times as likely to be a physicist than was a woman holding an advanced degree in 1980. In contrast, registered nursing was a far more common occupation for females of each education level than for males, with a woman holding an associate's degree more than 30 times as likely to be a registered nurse than was a man holding an associate's degree. In contrast, psychology was a similarly common occupation among men and women of each education level, with women a woman holding an advanced degree 1.7 times as likely to be a psychologist as was a man holding an advanced degree.

[Table I]

These differences in base-year characteristics mean that these three occupations vary substantially in their exposure to the changing gender composition of education groups shown in Figure I. These differences are illustrated in Figure II, which shows the percent female in each occupation predicted by the instrument and the actual percent female for these three occupations in each year from 1960 to 2010. As the fraction of degree holders who are female increases, the predicted percent female for each of these occupations increases, but the increase is much larger for psychology than for nursing or physics.

[Figure II]

The differences between the predicted changes in physics, nursing and psychology illustrate the identification of this paper. While these occupations differ from each other in many ways, the large predicted increase in the fraction of psychologists who are female, relative to the smaller predicted increases in the fraction of nurses and psychologists who are female, is not due to changes in psychology, in nursing or in physics. Instead, it is a consequence of the interaction of psychology's preexisting gender composition with broad changes in the role of women. As a result, these changes are far less likely to be confounded with changes in demand for these occupations, in the difficulty of work in these occupations, or in the returns to skills in these occupations than are actual changes in gender composition.

4.2.2: Additional Controls

The construction of this instrument produces two sources of endogeneity that require the introduction of additional controls. First, I control for the base-year education composition of each occupation interacted with year. Specifically, I include controls for $\gamma^{j}_{a,1980} * (\acute{t} = t)$, for each level of educational attainment *a* and year *t*, where $\gamma^{j}_{a,1980}$ is the fraction of workers in occupation j with education attainment a in 1980 and $\acute{t}$ is a particular census year. This controls for the time-varying effect of the base-year education composition on wage. This control is necessary because occupations that hire a larger fraction of workers from high education levels in the base year have larger implied increases in percent female. Because the returns to education have changed over time, this leads to a relationship between $\widetilde{f_{J,t}}$ and $W_{g,j,t}$ that is not related to the gender composition of the occupation.

Second, I control for the changes in labor supply across occupations brought about by the rise in women's relative education and labor force participation. Because this rise increased the supply of labor in occupations that predominantly hire highly educated women, it might be expected to lower wages in occupations with more female than male workers. This effect of rising women's education and labor force participation is theoretically distinct from the effect of gender composition because it depends on the number of workers available in the occupation, not whether those workers are male or female.

I disentangle the labor supply effect of the increase in women's relative education and workforce participation from its gender composition effect by constructing an index of "induced labor supply." This index estimates the fraction of the workforce that would work in occupation *j* in year *t* were men and women's education and labor-force participation decisions to change while their occupation decisions, conditional on education and labor-force participation, remained constant. Effectively, the induced labor supply index captures the mechanical relationship between the instrument Effectively, the induced labor supply index captures the mechanical relationship between the instrument Effectively, the induced labor supply index captures the mechanical relationship between the instrument for gender composition and labor supply. This labor supply index is defined below:

$$\widetilde{l_{J,t}} = log\left(\sum_A \gamma^{j}_{a,1980} * \left(\omega^{F}_{aj,1980} * k^{\bar{J}}_{f,a,t} + \omega^{M}_{aj,1980} * k^{\bar{J}}_{m,a,t}\right)\right) \qquad (4)$$

Where $k^{\bar{J}}_{f,a,t}$ is the fraction of the labor force at time *t* composed of females with educational attainment a, and $k^{\bar{J}}_{m,a,t}$ is the fraction of the labor force at time t composed of males with educational attainment a.

Like the gender composition index $\widetilde{f_{j,t}}$, this index holds the education composition of the occupation and the occupation choices of men and women in the workforce with each education level a fixed. However, rather than taking the ratio of the implied female workforce to the implied total workforce, it adds the implied female workforce to the implied male workforce. I take the log of this index in order to interpret coefficients on labor supply as elasticities. The log of the labor supply index is also more strongly correlated with wage than is a linear index normalized to 1980 labor supply, but the choice of functional form has small effects on the estimated effect of fraction female on wage.

It is important to note that this index is not designed to capture all changes in the labor supply of occupations. For instance, changes in the skill requirements, amenity values, or prestige attached to an occupation will affect labor supply, but will not be reflected in the index. Because some labor supply changes are mechanisms of effect, such a control would be undesirable.

The occupations with the largest induced changes in labor supply over the sample period are not the occupations with the largest induced change in fraction female. To show this, I depict the relationship between base-year gender composition by educational attainment and the instrument graphically in Figure III. This figure defines a summary statistic, "Weighted Average Female/Male Popularity by Education", which represents the average ratio of $\omega^F_{aj,1980}/\omega^M_{aj,1980}$, weighted by $\gamma^j_{a,1980}$. I compare changes in the instrument and to labor supply over the whole sample period in male-dominated occupations (occupations with a low weighted log gender ratio), in gender-balanced occupations (occupations with a weighted log gender ratio near 0) and in female-dominated occupations (occupations with a high weighted log gender ratio). This figure shows that occupations with the highest predicted changes in the instrument have balanced gender ratios, meaning that they are similarly popular among men and women of each education level. In contrast, changes in induced labor supply are highest in occupations with high ratios, indicating that they are disproportionately popular among female workers.

[Figure III]

This difference in distribution is crucial to disentangling the effect of fraction female from the effect of labor supply. A wage effect operating through labor supply should be greatest for female-dominated occupations, moderate for gender-balanced occupation, and lowest for male-dominated occupations. In contrast, a wage effect operating through fraction female should be greatest for gender-balanced occupations and low for both male-dominated and female-dominated occupations.

After including these controls, the differences in predicted changes in gender composition for physicists, nurses and psychologists remain significant. Figure IV shows the residuals of the instrumented and actual gender composition of the three example occupations when including educational attainment and labor supply controls, along with occupation and panel fixed effects and time-varying age and education controls. While nursing and physics are predicted to see declining percent female relative to the trend in gender composition of similarly educated occupations, psychology is predicted to see increasing percent female relative to similarly educated occupations. These predictions match the actual gender composition of these occupations.

[Figure IV]

4.2.3: Regressions of Interest:

I generate my main results by running a two-stage least squares analysis of equation 1, using the instrument defined in equation 4. The regression equations are below. Occupations are indexed by j, gender by g, and time lags (of 0, 10, and 20 years) by k. These regressions are of log mean wage $W_{g,j,t+k}$ onto fraction of the workforce that is female $\widetilde{f_{j,t}}$, log mean wage in year t for female and male workers $W_{f,j,t}$ and $W_{m,j,t}$, induced labor supply $\widetilde{l_{j,t}}$, time-varying average characteristics of female and male workers $X_{f,j,t}$ and $X_{m,j,t}$, occupation fixed-effects η_{jk} and year fixed-effects λ_{tk}.

First-Stage:

$$f_{j,t} = \beta_{1k}\widetilde{f_{j,t}} + \beta_{2kf}W_{f,j,t} + \beta_{2km}W_{m,j,t} + \beta_{3k}\widetilde{l_{j,t}} + \delta_{fk}X_{f,j,t} + \delta_{mk}X_{m,j,t} + \gamma_{jk} + \sigma_{tk} + \varepsilon_{j,t} \quad (5)$$

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$$W_{g,j,t+k} = \vartheta_{1k}\widetilde{f_{j,t}} + \vartheta_{2kf}W_{f,j,t} + \vartheta_{2km}W_{m,j,t} + \vartheta_{3k}\widetilde{l_{j,t}} + \vartheta_{4kf}X_{f,j,t} + \vartheta_{4km}X_{m,j,t} + \upsilon_{jk} + \varsigma_{t+k} + \pi_{g,j,t} \quad (6)$$

Structural Equation:

$$W_{g,j,t+k} = \theta_{1k}\widehat{f_{j,t}} + \theta_{2kf}W_{f,j,t} + \theta_{2km}W_{m,j,t} + \theta_{3k}\widetilde{l_{j,t}} + \theta_{4kf}X_{f,j,t} + \theta_{4km}X_{m,j,t} + \eta_{jk} + \lambda_{t+k} + \mu_{g,j,t} \quad (7)$$

4.2.4: Exogeneity of the Instrument:

As with all shift-share instruments, the exclusion restriction for this instrument depends on the panel structure of data and the presence of occupation and year fixed effects. Specifically, the instrument must be uncorrelated with the error term $\mu_{t,j}$ of structural equation (7)—i.e. there is no correlation between the instrument and present or future wage after controlling for year fixed-effects, occupation fixed-effects, education distribution at time *t*, age distribution at time *t*, induced labor supply, and education distribution in 1980 interacted with year, other than through $f_{j,t}$.

Goldsmith-Pinkham Sorkin and Swift (2018) (GPSS) note that this exclusion restriction holds in two cases. In the first case, the base-year fixed shares must be exogenous to changes in wage over time (equivalent to a common-trends assumption in a difference-in-difference estimation). In the second case, the time trends must be exogenous to dispersion in wages across occupation. GPSS formalize this argument for a traditional Bartik instrument, where the instrument is defined as the sum of industry shares multiplied by industry growth rates

. Because the instrument used in this paper interacts base year components (here the ratio of male to female

. Because the instrument used in this paper interacts base year components (here the ratio of male to female employment in occupation *j* given educational attainment *a*) nonlinearly with time trends (here the ratio of female to male workers with educational attainment *a*), it is not mathematically equivalent to IV GMM.

However, the exogeneity of the instrument still relies on the assumption that base-year employment ratios, $\Omega_{a,j,1980}$, affect changes in wages over time only through changes in gender composition. This assumption is equivalent to the common-trends assumption in a difference-in-difference framework. It assumes that, had the gender composition of occupations not changed, occupations that were gender-balanced, conditional on education, in 1980 would have experienced similar changes in wages as would occupations that were male-dominated or female-dominated.

This common-trends assumption will hold if base-year gender ratios reflect either idiosyncratic characteristics of occupations or persistent characteristics that do not have a changing relationship with wage. For instance, if women are more likely than men to work in caring occupations (Hirsch & Manzella, 2015), this might result in higher fractions of highly educated women choosing medical professions over engineering professions across the entire sample period. Likewise, idiosyncratic aspects of an occupation's culture or history can have longstanding effects on its gender composition. For instance, while the training, skills and work environments of dentists and veterinarians are very similar, dentistry was a far less popular occupation for female advanced degree holders than was veterinary medicine in every year of the data. As a result, dentists have a very low residual predicted growth rate in percent female over the sample period, while veterinarians have a fairly high residual growth rate in percent female. While a discussion of the relatively low number of female dentists by the American Dental Association highlighted a few challenges faced by women in dentistry, such as sexism by older male professors and practice owners and a lack of ergonomic dental tools designed for smaller hands (Solana, 2016), these factors likely reflect rather than determine the profession's gender composition.

The argument for the exogeneity of base-year gender ratios could be violated in two cases. First, gender ratios in 1980 could reflect persistent differences in occupations that have changing effects on wages over time. For example, among occupations where at least 50% of workers have bachelor's degrees, math requirements are lower in occupations that were relatively common among female workers in 1980 than in occupations that were relatively common among male workers (Supplementary Figure A.I). As a result, if the returns to math skills increased from 1960-2010, as suggested by Autor Levy and Murnane (2003), we would expect wages to increase in highly-educated male-dominated occupations and to decrease in highly-educated female-dominated occupations, regardless of changes in their gender compositions.

I investigate this potential source of bias by examining the relationship between potential confounders and 1980 gender ratios. I focus in particular on four skill and task indices identified by Autor Levy and Murnane (2003) and Deming (2017) as having large changes in returns over the past half-century. Using definitions from Deming (2017), I examine indices of mathematical/analytical skills, social skills, routine tasks and service-sector tasks measured in the 1998 o*net. These skill measures are described in detail in Deming (2017). As shown in Figure IV, while these skill and task indices do have relationships with the relative popularity of occupations by gender in 1980, none of them have peaks or troughs among occupations with mixed base-year gender composition. As a result, effects of these skills can be distinguished from effects of gender composition. As a robustness check, I include time-varying controls for the skill composition of occupations, and find that, as shown in Supplementary Table A.V, these controls have a very small effect on the estimated relationship between fraction female and wage.

[Figure V]

The argument for the exogeneity of base-year gender ratios could also be violated if gender ratios in 1980 reflect trends in characteristics like the factor share of labor or work amenities that are reflected in wage. For instance, if some occupations have allowed steadily more flexible working hours from 1960 to 2010, those occupations would be expected to have more flexible hours, and thus more female employees in 1980 than would comparable occupations that did not experience greater flexibility in working hours. As a result, they would expect to have declining wages over the sample period, as a consequence of steadily increasing flexibility.

I address this challenge by examining the relationship between the instrument and earlier realizations of log average wage. A valid instrument should affect wage in future periods but should not affect wage in past periods. As a result, future realizations of the instrument should be unrelated to present wage once controlling for the present realization of the instrument. I test this hypothesis by including $\widetilde{f_{J,t+10}}$ *and* $\widetilde{f_{J,t+20}}$ as controls in a regression of $W_{g,j,t}$ on $\widetilde{f_{J,t}}$. As shown in Supplementary Table A.VI, the estimated effect of future realizations of the instrument on contemporary wage are small relative to the standard error of the effect and relative to the estimated effect of the contemporary instrument. If the instrument were correlated with trends in wages, the future value of the instrument would have a negative relationship with wage of a similar magnitude to that of the contemporaneous value of the instrument.

As a second, more data-intensive approach, I simulate event studies by running the following regressions, for time-lags of -20, -10, 0, 10 and 20, indexed by *k*:

$$W_{g,j,t+k} = \alpha + \beta_1 \widetilde{f_{j,t}} + \beta_2 \widetilde{f_{j,t+k}} + \delta X_{j,t} + \gamma_j + \sigma_t + \varepsilon_{t,j} \tag{8}$$

The coefficient β_1 is presented in Figure VI for male wage and in Figure VII female wage. Because the instrument is highly autocorrelated, the event study on male wage is inconclusive. However, the event study on female wage is consistent with the identification strategy. There is no significant relationship between future realizations of the instrument and female wage, but there is a relationship between past realizations of the instrument and female wage. Were the instrument capturing the effect of a linear time trend, the coefficients should have a linear trend, rather than a discontinuity at t=0.

[Figure VI]

[Figure VII]

An alternative argument for validity of the instrument rests on exogeneity of the yearly percent female among workers of each educational attainment to dispersion in wages across occupations. For this argument to hold, trends in men's and women's education and labor supply decisions must originate For this argument to hold, trends in men's and women's education and labor supply decisions must originate from sources other than the wages paid in particular occupations. Goldin, Katz and Kuziemko (2006) identify several reasons behind the increase in women's educational attainment and labor force participation, such as relative increases in girls' aptitude test scores and likelihood to take math and science courses in high school, increasing fractions of girls expecting to work as adults, and delayed marriage. Currie and Morretti (2003) further note that women increased college attendance with the greater availability of coeducational schools, which increased in numbers considerably in the 1960's and 1970's (Goldin & Katz, 2011). Several papers have found that boys are more negatively affected than girls by adverse childhood environments, with Bertrand and Pan (2013) finding that negative home environment leads to greater behavioral problems for boys than for girls, and Autor et al (2016) finding that boys benefit more from high quality schools than do girls. This work suggests that the rise in single parent households and increases in economic inequality may contribute to the increase in women's relative education by rewarding girls' greater resilience. Likewise, the increase in women's relative labor force participation has been linked to several broad changes in the labor market, such as the elimination of policy barriers keeping married women out of teaching and clerical work in the 1950's (Goldin C. , 1988), changing expectations and social norms around women's labor force participation in the 1960's and 1970s, delayed marriage and child bearing, and a shift in women's identities that increased the value of career success (Goldin, Katz, & Kuziemko, 2006).

The exogeneity of the gender composition of educational attainment groups faces two challenges. First, exogeneity would be compromised if men or women choose whether or not to attend college or graduate school by considering the earnings of particular occupations. For example, if higher expected earnings for doctors led women to increase their graduate school attendance relative to men, this would result in covariance between the residual earnings of doctors and the gender composition of graduate school. Because I calculate the female fraction of workers with educational attainment *a* in year *t*, $f_{a,t}^{\bar{J}}$ by leaving out those in the aggregate occupation that includes j, women would need to enter graduate school on the basis of doctor's wages, but then not enter the medical field. This assumption would hold, however, if the gender composition of college graduates or postbaccalaureate graduates was associated with general increases or decreases in wages across occupations, because these general increases and decreases would be absorbed by the year fixed-effect. One piece of evidence against this concern is that controls for $W_{g,j,t}$ have no appreciable effect on the first-stage relationship between $\widetilde{f_{J,t}}$ and $f_{t,j}$ nor on the reduced-form relationship between $\widetilde{f_{J,t}}$ and $W_{g,j,t+10}$. If women and men choose levels of educational attainment based on expected future wages in particular occupations, controls for current period wage should matter. That they do not is evidence against this source of bias.

A more serious challenge to the exogeneity of gender composition of education groups with respect to relative wage is that the trends in the gender composition of education groups have been fairly smooth and linear over the past fifty years, particularly for bachelors' degree holders and advanced degree holders. As a result, the residual variation in the instrument after controlling for occupation linear time trends is insufficient for analysis. Identification from time trends in the gender composition of education groups could thus capture the effect of other trends that differentially affect the wages of occupations in the same time period. This concern is alleviated by the fact that future values of the instrument have a weak relationship with wage when controlling for current values of the instrument (Supplementary Table A.VI).

Section 5: Results

5.1: OLS Regression

Table II shows estimates of regression equations 2 and 3 on the decennial census from 1960 to 2010. Results are shown for regressions using controls for education and age composition of each occupation, as described in section 4.1, as well as for regressions that add induced labor supply and the time-varying effects of 1980 education composition as controls, as described in section 4.2.2.

[Table II]

These regression estimates suggest that there is a weak relationship or no relationship between changes in the gender composition of an occupation and the average wage of that occupation. However, as discussed in section 4.1.1, these results cannot be interpreted as a causal effect of gender composition on wages due to a multitude of non-causal factors that could explain the relationship between gender composition and future wage. In particular, the OLS results indicate that the rise in women's education and labor force attachment over the sample period is a significant source of positive bias in the relationship between changes in the gender composition of an occupation and changes in the wage of an occupation. In addition, measurement error is likely attenuating the OLS relationship between fraction female and wage—because most occupations have experienced relatively minor changes in gender composition during the sample period, particularly after controlling for year fixed-effects, residual variation is heavily influenced by sampling error.

This finding contrasts with a few papers in sociology that have found large, negative relationships between the gender composition of an occupation and future wages in that occupation using similar panel designs (Catanzarite, 2003), (Levanon & Allison, 2009)[8], but is in line with panel estimates using the CPS (England, Allison, & Wu, 2007). One explanation for the wide range of estimates in this literature comes from Tam (1997), who demonstrates that the presence of multiple endogenous relationships between gender composition and wage leads panel estimates to be highly sensitive to specification.

5.2: Instrumental Variables Estimates

Table III shows the results of the first-stage regression for the instrument, both for the full sample—of all occupations and years included in the definition of the instrument, and the lagged sample—of occupations for which $W_{g,j,t+10}$ is defined. The instrument $\widetilde{f_{J,t}}$ is a strong predictor of gender composition in both samples, with an f-statistic of 36 in the full sample and 37 in the lagged sample. The inclusion of wage controls $W_{f,t}$ and $W_{m,t}$ have no effect on first-stage estimates, while the inclusion of induced labor supply control $\widetilde{l_{J,t}}$ substantially increases the first-stage coefficient on the instrument. The negative effect on $\widetilde{l_{J,t}}$, and the positive effect of the inclusion of $\widetilde{l_{J,t}}$ on the estimated effect on $\widetilde{f_{J,t}}$ shows that male-dominated occupations had greater entry of female workers than implied by the instrument. As a result, previously female-dominated occupations have a lower percent female in later years than is implied by the instrument,

[8] Levanon, England and Allison (2009) uses a similar data source and panel framework as is used in this paper. Their results differ from the OLS estimates in this paper because their occupation definitions and functional form assumptions differ from this paper. In addition, in order to account for Nikkel bias Levanon, England and Allison fix the covariance between the residual and future values of the dependent variable to zero.

while previously male-dominated occupations have a higher percent female in later years than is implied by the instrument.

[Table III]

The two-stage least squares estimates shown in Table IV provide evidence that the fraction of workers who are female has large, negative effects on wages for males and females, both contemporaneously and over the following ten years. Estimated simultaneous effects are slightly higher for male than for female workers, with a 10 percentage-point increase in the fraction of female workers leading to an 8.2% decrease in average male wage and a 7.2% decrease in average female wage. Lagged effects are larger for female than for male workers, with a 10 percentage-point increase in the fraction of female workers leading to a 13.0% decrease in average female wage and an 8.7% decrease in average male wage. However, neither of these differences are statistically significant.

[Table IV]

There are a few notable characteristics of these findings that deserve some discussion. First, while the greater magnitude of lagged effects relative to simultaneous effects are not statistically significant, they could be a result of the fact that the proposed amenity and prestige mechanisms may take time to be fully realized. With respect to amenities, changes in the organization of occupations to allow flexible working hours, less competition, and other amenities of greater average value to female than to male workers may occur primarily through the formation of new firms. If so, responses to a change in gender composition may take time to occur. With respect to prestige, if changes in the prestige and advancement opportunities in an occupation lead the occupation to attract less ambitious workers, that effect will grow over time as workers change jobs and enter or exit the occupation.

Second, the lagged effect of fraction female on female wage may be larger than the lagged effect on male wage for two reasons. First, because occupations are not homogeneous, female workers may work in sub-occupations that see a larger average increase in percent female than do the sub-occupations of male workers. Estimates calculated using occupations aggregated to the sub-header level (Supplementary Table A.VII, described in section 5.3) support this hypothesis—estimates using aggregated occupations show larger differences between the effect on male wage and on female wage than do estimates using disaggregated occupations. Second, male and female workers may be complements in production (Giorgi, Paccagnella, & Pellizzari, 2013). In this case, the presence of a large number of female workers would increase the productivity of male workers, mitigating negative effects on wages. This might be the case if, for instance, male workers are more likely to hold managerial roles than are female workers.

Finally, these results do not appear to be highly sensitive to the inclusion of wage and labor supply controls. Controlling for current-period wage does not substantially alter the estimated effect on either male or female wage—the increase in estimated effect is less than half a standard error for males and females in the absence of a control for current-period wages. This provides evidence that these results are not driven by correlated trends in occupation wages and the gender composition of education groups. The small difference in estimated effect also reduces concern for Nikkel Bias arising out of the inclusion of a lagged dependent variable in a panel regression (Kiviet, 1995). Likewise, controlling separately for implied male and female labor supply, $\widetilde{l_{f,J,t}}$ and $\widetilde{l_{m,J,t}}$, rather than controlling for overall implied labor supply $\widetilde{l_{J,t}}$ causes a reduction in the estimated effect on male and female wage of less than a standard error.

In order to capture long-term effects of gender composition, I examine the effect of percent female in time *t* on log mean wage at time *t+20* for men and women, shown in Table V. I find that while wages remain substantially lower for males and females, the effects are smaller 20 years after the change than they are 10 years after the change. A 10 percentage point increase in female share in time *t* lead to a 8.6% decrease in average female wage by time *t+20* and a 4.5% decrease in average male wage by time *t+20*. This reduced negative effect of gender composition on wage could result from firms learning how to operate efficiently with forms of organization that allow for more flexible hours and less competitive work environments. It could also reflect fade-out of negative effects of fraction female on prestige.

[Table V]

One puzzle in interpreting these results is why the estimated causal effect of fraction female on wage is so large. While female-dominated occupations pay less than do male-dominated occupations, the cross-sectional relationship between log average wage and fraction female is smaller than the estimated causal effect of fraction female on average wage. When controlling for the age and educational composition of occupations, the cross-sectional effect of a 10 percentage point increase of percent female in 2010 was 4.2% for female wage and 4.8% on male wage (Supplementary Table A.I). Thus, extrapolating the local average treatment effect estimated in this paper to differences in cross-sectional gender compositions would imply that the causal impact of gender composition explains or more than explains differences in pay between male-dominated and female-dominated occupations. However, there are several explanations for this finding that allow intrinsic characteristics of occupations (any characteristic not affected by the gender composition of the occupation) to have some effect on wages.

First, because female workers are more educated on average than are male workers, cross-sectional differences in wage between male and female workers are greater when including education controls. Because the education controls estimated from the decennial census are fairly broad (for instance, aggregating masters, professional and doctoral degrees), it is possible that the cross-sectional differences in wage would be larger if including a richer set of education controls. Because the estimations of the causal impact of fraction female on wage include occupation fixed-effects, they effectively control for arbitrarily detailed differences in education by occupation (averaged across time-periods). As a result, the estimated causal effect of changes in gender composition on wages may be a smaller fraction of a more fine-grained measure of the cross-sectional relationship between the gender composition and wage of an occupation.

Second, I find evidence that the effects of a change in gender composition are not entirely persistent. The estimated effect of a change in gender composition at time t is larger on wage in time $t+10$ then on wage at time $t+20$, with effects on female wage falling from 13% to 9%, and effects on male wage falling from 9% to 5%. This fade-out is theoretically reasonable—for instance, if occupations begin to reorganize in order to reduce the returns to long hours, the reorganization may be costlier to productivity in the short-run than in the long-run. In addition, because the instrument features significant intertemporal autocorrelation, lagged estimates are likely biased in the direction of short-term estimates (Jaeger, Ruist, & Stuhler, 2018), so actual effects could fade out to a greater extent than estimated. As a result, the long-run

Formatted: Font: Times New Roman, 11 pt, Do not check spelling or grammar equilibrium effect of fraction female on wage could be quite a bit smaller than the estimated 20-year effect. In this case, the negative wage effects of a high percent female in occupations that have been female-dominated for a long time, like nursing, may be very small.

5.3: Robustness Checks

5.2.1: Alternative Instruments:

I define two alternative instruments that may be less exposed to confounding time trends than is the primary instrument. While these instruments are too weak to provide consistent, unbiased estimates, they produce results that are in line with the results from the main specification.

I construct a first alternative instrument that holds fixed workers' college major decisions, rather than their occupation decisions. While a future worker's college major decision is influenced by their expected future occupation, the college major decision is made at a degree of remove from the labor market, and at a time when future workers have less knowledge about the earnings and job amenities of various occupations. As a result, gender differences by major choice may be less likely to reflect changes over time in the characteristics of the occupations they feed into than do gender differences in the occupations themselves. As shown in Table VI, men and women choose different college majors, conditional on their level of educational attainment[9]. Because of these persistent differences in the college major choices of women and men, the effect of women's increased educational attainment on gender composition is greater for occupations that predominantly hire from gender-balanced college majors, like Biological/Life Science rather than from male or female-dominated majors, like Education or Physical Science/Engineering.

[Table VI]

[9] The majors described in Table VI are aggregates of the college majors listed in the American Community survey. The majors included in each aggregate are presented in Supplementary Table A.XIV.

This instrument estimates the fraction of available workers in each occupation who would be female if the occupation hired the same proportion of workers from each education level and college major as it did in 2010, if male and female students had the same propensity to choose each college major as they did in 2010, but if the fraction female of each educational attainment group varied over time. I define the instrument as:

$$\widetilde{f_{t,J}}^{M} = \sum_{A,M} \gamma^{j}_{am,2010} * \left(\frac{\omega^{F}_{am,2010} * f^{\bar{J}}_{a,t}}{\omega^{F}_{am,2010} * f^{\bar{J}}_{a,t} + \omega^{M}_{am,2010} * \left(1 - f^{\bar{J}}_{a,t}\right)}\right) \quad (10)$$

Where $\omega^{F}_{am,2010}$ is the fraction of females with educational attainment *a* choosing major *m* in 2010, $\omega^{M}_{am,2010}$ is the fraction of males with educational attainment *a* choosing major m in 2010, $f^{\bar{J}}_{a,t}$ is the female fraction of workers with educational attainment *a* in year *t*, excluding those employed in occupation *j*, and $\gamma^{j}_{am,2010}$ is the fraction of workers in occupation j with education level *a* and major *m*, in 2010. Because variation in this instrument depends on college major, which is available only for college graduates, I include only occupations where at least 25% of workers had a college degree or higher in 2010 in estimates using this instrument.

I construct a second alternative instrument by calculating the instrument defined in equation 4 using 1970 as a base year, rather than 1980. I exclude 1960 data from the estimates using this instrument, so that 1970 is both the base year and the initial year. I use 1970 as a base year, rather than 1960, because the female fraction of most professional occupations in 1960 was too low to generate reasonable estimates. This instrument is estimated as:

$$\widetilde{f_{t,J}}^{1970} = \sum_{A} \gamma^{j}_{a,1970} * \left(\frac{\omega^{F}_{aj,1970} * f^{\bar{J}}_{a,t}}{\omega^{F}_{aj,1970} * f^{\bar{J}}_{a,t} + \omega^{M}_{aj,1970} * \left(1 - f^{\bar{J}}_{a,t}\right)}\right) \quad (9)$$

I construct this instrument to address the possibility that the main result is confounded with trends in characteristics of occupations that are related both to wages and to the relative likelihood that males and females work in an occupation. For instance, suppose that social science and therapeutic occupations experienced decreasing returns to long hours, while returns to long hours remained high in physical science and engineering occupations. This could lead educated women to choose social science and therapeutic occupations at increasing rates relative to physical science and engineering occupations, leading to a correlation between the conditional gender ratio of occupations in 1980 and the trend in returns to hours over the sample period.

Several pieces of evidence are inconsistent with such a trend being present through the full sample period. Because such a trend would result in the shift-share instrument not fully accounting for changes in occupational sorting between men and women, it would bias a shift-share instrument in the direction of OLS results, which do not account at all for changes in sorting. Instead, I find instrumental variables estimates that are far larger than the OLS estimates. Additionally, as shown in Supplementary Table A.VI, future values of the instrument are not correlated to wage, controlling for the current value of the instrument. However, such a trend may be present for some part of the sample period. In particular, changes to occupations from computerization or automation may only begin to affect gender sorting and wages from 1980 onward. A 1970 base year eliminates trends that only begin to affect sorting by gender in 1980.

These two alternative instruments are correlated with the primary instrument, but the residual variation in the alternative instruments differs substantially from the residual variation in the primary instrument. Supplementary Table A.VIII shows the correlations between the residual of each instrument after controlling for occupation and year fixed effects, current-period education and age, time-varying base-year education and induced labor supply.

Table VII shows the estimated effect of fraction female on wage, using these robustness instruments. These estimates are run on restricted samples. Estimates using the college major instrument, exclude occupations where fewer than 25% of workers were college graduates in 2010, as well as occupations that were not included in the 2010 occupation codes or do not include an induced labor supply control. Estimates using a 1970 base year exclude 1960, and use an induced labor supply control, $\widetilde{\widehat{l_{J,t}}}^{1970}$, that is defined using a 1970 base year.

[Table VII]

These robustness checks produce fairly noisy estimates, but the results are broadly consistent with the results of the primary analysis. Estimates using the college major instrument indicate that a 10 percentage point increase in fraction female leads to a 21% decline in contemporaneous female wages that persists over 10 years. The estimated effect on male wages is noisy and insignificant, with a 10 percentage point increase in the fraction female leading to a 0.5% increase in contemporaneous male wage and a 1.6% decrease in male wage over 10 years. Estimates using a 1970 base year are also broadly consistent with those using the primary instrument. Using a 1970 base year, a 10 percentage point increase in fraction female leads to an estimated contemporaneous 6% decline in average female wage and a 14% decline in average male wage. Over 10 years, the effect shifts to a 14% decline in average female wage and a 9% decline in average male wage. Because the standard errors on these estimates are very wide, the difference between these results and the primary results may be statistical noise. This is especially likely because estimates from each robustness instrument are not consistently larger or smaller than the primary results.

I present seven additional robustness checks, with results in Supplementary Tables A.VII-A.XII. Each is described in turn below.

Estimate results using College Graduates only (Supplementary Table A.IX)

I estimate the effect of changes in gender composition on the wages of college graduates as an additional test for bias resulting from changing returns to education. Because college-educated workers saw increasing relative wages over this period, changing returns to education could be confounded with changes in the gender composition of education groups. As show in Table A.IX, the estimated effect of a change in fraction female on log male and female wage is similar for college graduates and for all workers. An increase in the fraction female of 10 percentage points leads to an 8% decrease in female wages and a 9% decrease in male wages contemporaneously, and a 19% decrease in female wages and a 10% decrease in male wages over 10 years.

Estimate wage effect on "Newcomer" workers aged 20-35 (Supplementary Table A.X)

I examine the wages of young workers in order to determine whether the estimated effects of fraction female on log average wage reflect complementarities between young and old workers. In particular, a negative effect on the wage of older workers from an increased fraction female among younger workers could be caused by young male workers enhancing the productivity of older workers. Because the estimated effect of fraction female on the average wage of young female workers (Table A.X) is similar to the effect on older female workers, this does not appear to be the case. However, the estimated effect of fraction female on the average wage of young male workers is less negative than is the estimated effect for older male workers, with a 10 percentage point increase in the fraction female leading to a 5% decline in young male wage and a 9% decline in older male wage over 10 years. This may be a consequence of competition between young workers—young male workers may be more likely to be promoted when a larger fraction of the young workforce is female, either due to gender bias or due to differences in preferences and productivity. Additional evidence for the effect of competition comes from the greater estimated labor supply elasticity for young male and female workers than for older workers, with a 10% increase in the labor supply of young workers leading to a 2.5% decrease in the wages of young female workers and a 1.5% decrease in the wages of young male workers, measured contemporaneously. The same increase in labor supply leads to a 0.4% decrease in the wages of older female workers and a 0.3% decrease in the wages of older male workers.

Estimate results without labor supply controls and without time-varying base-year education controls (Supplementary Table A.XI).

I examine the effect of fraction female on log wage, removing labor supply and time-varying education controls from the analysis. Time-varying education controls are included in the main analysis because the instrument is positively correlated with the education level of the occupation in the base year. While time *t* education controls are included in all models, it is possible that occupations that had higher measured education requirements in 1980 also have higher unmeasured education requirements in all years, conditional on measured education requirements in time *t*. However, because the inclusion of time-varying education controls absorbs a great deal of variation in the instrument, it is worthwhile to see whether the results are robust to the exclusion of this control. While the exclusion of 1980 base-year controls reduces the magnitude of the estimated effect of fraction female on average wage, particularly for male wages, estimated effects remain substantial and negative. A reduction in the estimated effect of fraction female on average wages is consistent with increasing returns to education.

Estimate results using aggregate occupation definitions (Supplementary Table A.V)

Changes in the classification of occupations over the period 1960 to 2010 add noise to the estimated effect of fraction female on average wage. These reclassifications could also bias the estimated relationship if high-wage jobs are likely to be classified differently than similar, predominately female jobs. To address this, I estimate the relationship between fraction female and wage using 1990 occupation definitions, aggregated to the sub-header level. The three-digit Standard Occupation Codes used in the decennial census are published with codes grouped under thematic sub-headings. For example, Physicians, Dentists, Veterinarians, Optometrists, Podiatrists, and Health diagnosing practitioners, n.e.c. are all included in the sub-heading "Health Diagnosing Occupations." I use sub-heading aggregation, rather than aggregating Standard Occupation Codes to the two-digit level, because occupations grouped by sub-heading are more consistent with each other than are occupations with the same first and second digit, and because these occupations were intentionally categorized together in the construction of the occupation codes.

As shown in Supplementary Table A.V, a 10 percentage point increase in the fraction female of an aggregate occupation leads to an estimated 8% decline in average female wage and a 6% decline in average male wage, measured contemporaneously. Over 10 years, the effect grows to a 14% decline in average female wage and a 5% decline in log male wage. While these results are broadly consistent with the estimated effects on 3-digit occupations, effects on male wages are smaller than on female wages. This may be a consequence of male workers having smaller changes in their specific occupations than is implied by the gender composition of aggregated occupations.

Estimate results including control for employment as a share of the total workforce (Supplementary Table A.XII)

I include controls for the log of current period employment (measured as the fraction of all workers who work in occupation j at time t). I do this to address concerns that the labor supply index, defined in equation 4, does not fully account for changes in labor supply of occupations induced by women's increased education and workforce participation. A possibility of particular concern is that the occupations with the greatest increase in fraction female also saw the greatest actual increase in labor supply. This would occur if, among occupations that were open to women, women chose gender-balanced occupations in greater numbers as their share of the workforce increased. Despite these concerns, I exclude actual employment from my main specification for two reasons. First, because changes in selection into an occupation are a causal channel, employment controls may absorb some of the causal impact of gender composition on wage. Second, because occupation reclassifications have a greater effect on employment than on wage, employment data from the census is noisy.[10] As shown in Table A.XII, employment controls do not change the estimated effect of fraction female on wage.

Estimate Long-Difference Results (Supplementary Table A.XIII)

As a final test of the effect of gender composition on wage, I estimate the relationship between expected change in percent female over the period from 1960 to 2010 and the change in the average wage of male and female workers in the same period. The long difference approach is valuable for two reasons. First, it does not take a stance on the timing of the effect of fraction female on log wage. Second, it accepts the relative linearity of trends in the gender composition of education groups by treating the entire period as a single shock.

I estimate long-difference models using first-difference estimation for occupations that were included in the 1960 and 2010 census. I calculate Δf_j, $\Delta \widetilde{f_j}$, $\Delta \widetilde{l_{j,t}}$ and $\Delta W_{g,j}$ as the difference between the 2010 and 1960 values of fraction female, instrumented fraction female, induced labor supply and wage respectively. I then estimate the following two-stage least-squares equations, with results shown in Supplementary Table A.XIII:

First-Stage:

$$\Delta f_j = \beta_1 * \Delta \widetilde{f_j} + \beta_2 * \Delta \widetilde{l_j} + \beta_3 \Delta X + \varepsilon_j \qquad (11)$$

[10] Employment is more affected by occupation reclassifications than is wage or gender composition because employment is affected by all grouping or splitting of occupations, even if the grouped and split occupations are very similar. In contrast, grouping two similar occupations (that have similar gender compositions and wages) together will have a modest effect on measured average percent female and wage.

Reduced-Form:

$$\Delta W_{g,j} = \beta_1 * \Delta\widetilde{f_j} + \beta_2 * \Delta\widetilde{l_j} + \beta_3 \Delta X + \varepsilon_j \tag{12}$$

As shown in Table A.XIII, the results of the long-difference regression are noisy and are not statistically significant at conventional levels, but are broadly consistent with the main estimates of the effect of fraction female on wage. A 10 percentage-point increase in fraction female leads to an estimated decline over the sample period of 3.6% for men and 9.1% for women. These estimates are in line with the 20-year estimated effects of 8% for women and 5% for men.

Section 6: Mechanisms

As discussed in section 2, while there are several mechanisms by which the gender composition of an occupation can affect wages paid in that occupation, the two sets of mechanisms most likely responsible for these results are changes in workplace amenities and changes in prestige. The amenities mechanism operates through differences in the preferences of male and female workers. As an occupation sees an increased percentage of female workers, the hours requirements, competitive pay structures, and other job characteristics adjust to reflect the preferences of female workers. The prestige mechanism operates through differences in the perceived abilities of male and female workers. As an occupation sees an increased percentage of female workers, male and female workers in the occupation are perceived to be of lower ability or lower quality by outside employers, causing reduced advancement opportunities and making the occupation less attractive to high ability/high ambition workers. These two mechanisms are not mutually exclusive, and may in fact be mutually reinforcing.

I perform two tests of the amenities mechanism. First, I examine the effect of a change in fraction female on the average number of hours worked per week by men and women. If women pay higher costs for long working hours than do men, increased representation of women in an occupation will create incentives for firms to organize work in a way to reduce the cost of short working hours. As a result, the average number of hours worked should decrease for both male and female employees. I test this hypothesis by estimating a two-stage least-squares regression of average hours worked on fraction female. These results are presented in Table VIII.

[Table VIII]

As shown in Table VIII, there is a negative relationship between fraction female and hours worked, consistent with an amenities channel. The estimated effect of gender composition is similar in magnitude to the difference in average hours worked between male and female workers. However, the difference is not statistically significant when controlling for labor supply.

I perform a second test of amenities by estimating the elasticity of earnings with respect to hours worked per week for each occupation and year, following Goldin (2014). In each census year *t*, I perform the following regression at the individual level:

$$logearn_i = \alpha_{j,t} + \beta_{j,t} loghours_{i,j} + \gamma_{j,t} fem_{i,j} + \delta_t X_i + \varepsilon_i \qquad (13)$$

Where $logearn_i$ gives log annual earnings for worker i, $loghours_{i,j}$ gives the log number of hours per week reported by worker *i*, $fem_{i,j}$ indicates whether worker *i* is female, and X_i gives a set of individual controls that includes log weeks worked per year, a quartic of age, and years of education. $\beta_{j,t}$ measures the elasticity of earnings with respect to hours worked for full-time workers in occupation *j* at time *t*. This regression is run on a sample of full-time workers (working at least 35 hours per week and 40 weeks per year) between the ages of 25 and 65. By estimating the regression equation in each year, I create a measure of the elasticity of earnings with respect to hours at the year by occupation level. Because full-time workers who work long hours may differ from those who work short hours within an occupation, the estimated elasticity of earnings with respect to hours, $\beta_{j,t}$ is biased by selection.

I measure the effect of fraction female on returns to hours worked by regressing fraction female on $\beta_{j,t}$, using the regression specification described in equation 6. In order to account for measurement error, I weight estimates by the inverse standard error of $\beta_{j,t}$. As shown in Table IX, changes in fraction female is positively associated with returns to hours worked.

[Table IX]

The effects of fraction female on returns to hours worked are modest. A 10 percentage point increase in the fraction female would increase returns to hours worked by 0.114, implying that a 10% increase in hours worked would increase earnings by 0.11% more than it would have prior to the change in gender composition. This effect fades out over time, shrinking from an estimated contemporaneous effect of 0.91 to a 20-year effect of 0.45.

This result is consistent with a higher fraction female leading to greater demand for short hours, followed by a reorganization of firms to reduce the costs of short hours. Goldin (2014) argues that increases in the number of women in an occupation should increase the returns to long hours in the short run. If demand for the amenity of short working hours increases and the supply of short working hours is upward-sloping, the cost of short working hours should increase. At the same time, greater demand for short working hours increases returns for business models that make short working hours less costly. The decreasing effect on returns to hours worked over 20 years is consistent with this analysis—as firms learn how to reduce the cost of short working hours, the earnings differences between workers working long hours and workers working short hours shifts toward its previous level, even as more workers reduce their hours. However, it is also important to note that the elasticity of earnings with respect to hours may also be biased by selection into long hours of work. Because increases in fraction female are associated with decreases in hours worked, the workers who continue to work long hours may be higher wage than those who reduce their hours.

Finally, I test the effect of changes in the female fraction of an occupation's workforce on prestige. I measure changes in an occupation's prestige by taking the difference between the prestige of an occupation measured by the 1964 Hodge-Siegel-Rossi survey (Siegel, 1971) and the prestige of the occupation measured by the 1989 General Social Survey (Nakao & Treas, 1994). Each survey measures the prestige of an occupation by asking respondents to place each occupation on a ladder with ten rungs. Scores are then calculated as the weighted average of the score given by each respondent, and placed on a scale from 1-100.

I estimate the effect of fraction female on prestige by performing a two-stage least-squares regression of the difference in measured prestige from 1964-1989 on changes in the predicted female fraction of the occupation from 1970-1990. The results are shown in Table X.

[Table X]

Change in prestige has a weak negative association with the change in the fraction of female workers in an occupation. When controlling for time-varying education effects, a 10 percentage-point increase in the fraction female of an occupation is associated with a 1.6 point decrease in measured prestige. Without time-varying education controls, the effect is larger, with a 10 percentage-point increase in fraction female leading to a 3.3 point decline in measured prestige. This provides suggestive evidence that changes in the gender composition of an occupation might change the perceptions of that occupation among the public.

Section 7: DISCUSSION

In this paper, I find causal evidence on the effect of gender composition of an occupation on the wages of that occupation. My findings indicate that the effect of an increase in the percent female of an occupation on the wages for men and women are large and sustained, with a 10 percentage-point increase in the female share of an occupation leading to an 8% decline in average male wage and a 7% decline in average female wage. That effect grows over ten years to a 9% decline in average male wage and a 13% decline in average female wage, and shrinks over 20 years to a 5% decline in average male wage and a 9% decline in average female wage. These effects are large, and if the effect of gender composition on wage is homogeneous would explain or more than explain the cross-sectional relationship between gender composition and wage. While these estimates are imprecise, even lower-bound estimates are large relative to the absolute difference in average pay between male-dominated and female-dominated occupations.

These findings have several implications for policy and future research. Most importantly, these findings suggest that changes to occupations in response to the gender composition of the occupation's workforce play a crucial role in producing the observed discrepancy between the wages of female-dominated and male-dominated occupations with similar education requirements. As a result of lower pay in occupations with more women, preferences and skills more common among female workers may earn lower returns in the labor market than they would were gender a less salient factor in the labor market.

These findings also suggest that many of the observable differences in the characteristics of male and female dominated jobs may be consequences of gender composition, including differences in prestige, in returns to hours, and in ownership structures. This in turns suggests that correlations between observed characteristics of occupations that are related to gender composition and wages are likely to reflect the effect of gender composition on wage, and cannot be taken as measures of the direct effect on the observable characteristics themselves.

Other implications of this work depend on deeper exploration of the mechanisms driving the result. If the decline in wages in occupations that see rising female shares of the labor force is caused by the provision of amenities valued by women, this finding may auger well for gender integration of occupations. An amenities mechanism implies that occupations with work and pay structures that do not appear conducive to female work, such as business and management (Bertrand, Goldin, & Katz, 2010) may become more conducive to female work in the presence of a larger female workforce. Likewise, it implies that were men and women to continue to become more similar in their balancing of work and non-work responsibilities (Goldin C. , 2014), differences in the amenities provided in male and female-dominated occupations, and thus the earnings of those occupations, could also become more similar.

On the other hand, if the decline in wages in occupations with rising female shares of the labor force is caused by declines in prestige for those occupations, this finding would suggest that wage gaps may persist even in the face of continued integration of currently male-dominated occupations. A prestige mechanism implies that integrating a high-paying male-dominated occupation will likely cause declines in the wages paid in that occupation for male and female workers, due to some combination of highly skilled workers choosing other occupations, demand for the occupation falling, and a reallocation of high-skill tasks away from the occupation. As a result, the entry of substantial numbers of female workers into a particular highly paid male-dominated occupation may not generate as large an increase in earnings for those workers as would otherwise be expected. Future work should look more deeply into these mechanisms by examining changes in the organization of firms and the allocation of tasks among workers in response to changes in gender composition.

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Data Appendix

I use the decennial census to measure the mean wage and demographic composition of each occupation in the years from 1960-2000, and the American Community Survey in 2010. I define key variables in the following way:

Occupation: I define occupation using a modification of the 1990 occupation codes constructed by IPUMS. These codes are constructed using crosswalks to earlier and later occupation coding schemes, as described by Myer and Osborne (2005). Each crosswalk is constructed by assigning a group of workers from a census year both to an occupation in the previous scheme and an occupation in the new scheme. The new occupation codes are then linked to the old occupation codes by choosing the old code most frequently selected by those with each new code. Because the Standard Occupation Codes underwent a significant reclassification in 1970 and in 2000, the 1990 occupation scheme was chosen to avoid transforming any year's data with two major reclassifications.

I deviate from the 1990 harmonized codes in the 2000 and 2010 census by using the gender-specific occupation classifications provided by Blau, Brummund and Liu (2013). Blau et. al. observed that because the most prevalent 1990 codes for each 2000 code differ for male and female workers, gender segregation of occupations appears to decrease in 2000 when measured with harmonized 1990 occupation codes. To address this, Blau et. al. define a gender-specific crosswalk between the 1990 and 2000 codes by choosing the most prevalent 1990 occupation separately for women and men for each 2000 occupation code.

Wage: I calculate wage in the decennial census by dividing wage and salary income by estimated annual hours worked. The census asks respondents the number of hours they worked in the past week and the number of weeks they worked in the last year. Respondents choose weeks worked per year as intervals. I use the midpoint of each interval as the respondent's weeks worked. The census changed its question on hours worked after the 1990 census. From 1940-1990, the census asked respondents how many hours they worked last week, and provided their responses as intervals: 1-14, 15-29, 30-34, 35-39, 40, 41-48, 49-59, 60+. From 1980-2010, they asked respondents how many hours per week they usually work, with allowed responses ranging from 0-99. I estimated hours worked per year from 1960-1990 by assigning each individual to the midpoint of each interval, and multiplying hours per week by weeks per year. I estimated hours worked per year in 2000 and 2010 by multiplying usual hours worked per week by weeks worked per year.

Education: I divide respondents into five categories by educational attainment, using the detailed educational attainment variable. For years prior to 1990, the categories were defined as follows:

1) Less than high school: Respondents with no schooling completed, less than 12 grade completed, or 12 grade completed with no high school diploma.
2) High school graduate: High school graduate or GED recipient, less than 2 years of college and no college degree.
3) Some College: 2-3 years of college
4) College graduate: 4-5 years of college
5) 6+ years of college

From 1990 onward, they were defined as:

1) Less than high school: Respondents with no schooling completed, less than 12 grade completed, or 12 grade completed with no high school diploma.
2) High school graduate: High school graduate or GED recipient, less than 2 years of college and no college degree.
3) Some College: 2-4 years of college, no bachelor's degree
4) College graduate: 4+ years of college, no master's, graduate or professional degree
5) Master's, graduate or professional degree

These category definitions follow Jaeger (1997), who finds that these definitions minimize the distance between 1980 and 1990 education definitions. While there are no discontinuities in the gender composition of the some college, college graduate, and master's degree categories from 1980 to 1990, the percent of the workforce in each of these categories does change, with fewer advanced degree holders identified in the 1990 census than in the 1980 census. This is of potential importance to the construction of the labor supply control discussed in section 3.1, but does not appear to affect the residual of labor supply once controlling for year fixed effects.

American Community Survey (2009-2014)

Beginning in 2009, the American Community Survey asked respondents with a college degree to provide their college major, coded into 38 categories. Respondents were allowed to list up to three majors. Because the data demands of calculating the fraction of workers in each occupation with a particular education type are high, I collapsed the original 38 categories into ten broader categories, listed in Supplementary Table I. Occupation was determined from the American Community Survey using the Blau modification of the 1990 occupation codes described for the decennial census.

Panel Study of Income Dynamics:

The Panel Study of Income Dynamics (PSID) surveys a nationally representative sample of 18,000 respondents selected in 1968 and their descendants. Surveys are annual from 1968-1997, and semi-annual from 1999-2015. I use the Panel Study of Income Dynamics to look at the effects of changing gender composition on entry to and exit from occupations. The primary variables used in the PSID are occupation and wage income.

Occupation: The PSID uses 1970 occupation codes from 1968 to 1999, after which it switches to 2000 occupation codes. I crosswalk these codes to the modified 1990 occupation codes described above, and then aggregated based on aggregated occupations listed in the 1990 SOC codebook. A worker is determined to have changed occupation when their occupation differs from their occupation in the previous year.

Wage and Salary Income: Analyses on the PSID are conducted on Wage and Salary Income, rather than calculated hourly wage. This was done because the definition of wage and salary income is more consistent across years of the PSID than are definitions of hours and weeks worked, or direct questions on hourly pay. Wage and Salary income includes the labor component of business and farm income, wage income, bonuses, overtime and commissions, and income from professional practice.

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Tables

TABLE I: BASE-YEAR FIXED COMPONENTS FOR THREE EXAMPLE OCCUPATIONS:

	%of Workers With Ed	Workers / 10,000 Females	Workers /10,000 Males	Female / Male Ratio
Physicists and astronomers				
1: No Degree	0.53	0.03	0.11	0.31
2: HS Degree	3.18	0.07	0.42	0.16
3: Associates Degree	7.56	0.30	1.55	0.19
4: College Degree	17.90	0.57	5.03	0.11
5: Advanced Degree	70.82	3.18	36.64	0.09
Psychologists				
1: No Degree	1.22	1.02	0.39	2.62
2: HS Degree	2.55	1.05	0.51	2.07
3: Associates Degree	4.55	2.68	2.45	1.10
4: College Degree	15.21	14.66	8.87	1.65
5: Advanced Degree	76.47	177.96	104.55	1.70
Registered nurses				
1: No Degree	1.27	16.39	1.13	14.53
2: HS Degree	7.36	48.57	2.43	20.02
3: Associates Degree	55.09	739.75	24.34	30.39
4: College Degree	30.03	568.05	26.69	21.28
5: Advanced Degree	6.26	345.46	24.35	14.19

*Notes: % of workers with ed gives the percent of workers in occupation j with degree a in 1980, workers/10,000 Females gives the percent of women with degree a working in occupation j, and workers/10,000 males gives the percent of men with degree a working in occupation j. Female/Male Ratio gives (Workers/10,000 Females)/(Workers/10,000 Males)

TABLE II: OLS ESTIMATES

	Log Female Wage				Log Male Wage			
	(t)	(t)	(t+10)	(t+10)	(t)	(t)	(t+10)	(t+10)
Fraction Female	-0.04 (0.08)	-0.1 (0.08)	0.02 (0.07)	-0.05 (0.07)	-0.01 (0.07)	-0.06 (0.07)	0.11 (0.07)	0.07 (0.08)
Female Wage			0.03 (0.07)	0.02 (0.07)			0.09*** (0.04)	0.08** (0.03)
Male Wage			0.25*** (0.08)	0.26*** (0.08)			0.14** (0.06)	0.15** (0.07)
Labor Supply		-0.01 (0.09)		-0.16* (0.08)		0.01 (0.06)		0.02 (0.07)
Base-year Controls		X				X		X
Sample Size	1817	1817	1457	1457	1817	1817	1467	1467

*Note: Each column reports results from an estimate of equation (1) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education, and occupation and year fixed- effects as described in the text. Base-Year Controls are the share of the workforce in each education category in 1980, interacted with year. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE III: FIRST-STAGE REGRESSION

	Fraction Female (Full sample)			Fraction Female (Lagged Sample)		
	(3)	(2)	(1)	(6)	(5)	(4)
Instrument	0.51***	0.72***	0.72***	0.58***	0.86***	0.85***
	(0.11)	(0.12)	(0.12)	(0.13)	(0.14)	(0.14)
Labor Supply		-0.14***	-0.14***		-0.18***	-0.17***
		(0.03)	(0.03)		(0.04)	(0.04)
Male Wage			0.00			0.01
			(0.03)			(0.03)
Female Wage			-0.01			-0.02
			(0.01)			(0.01)
Number of Observations	1817	1817	1817	1457	1457	1457
Number of Occuaptions	380	380	380	379	379	379

*Note: Each column reports results from an estimate of equation (5) in the paper, with Log fraction of workers aged 22-65 who are female as the dependent variable, for occupations that have measured wages for male or female workers. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE IV: TWO-STAGE LEAST SQUARES ESTIMATES OF PERCENT FEMALE ON LOG MEAN WAGES

	Females					Males				
	Log Wage (t)		Log Wage (t+10)			Log Wage (t)		Log Wage (t+10)		
2SLS:	(1)	(2)	(3)	(4)	(6)	(7)	(8)	(9)	(10)	(12)
Fraction Female	-0.86*	-0.72**	-2.09***	-1.47***	-1.30***	-0.92***	-0.82***	-1.12***	-1.01***	-0.87***
	(0.59)	(0.39)	(0.71)	(0.42)	(0.42)	(0.34)	(0.28)	(0.40)	(0.30)	(0.27)
Labor Supply		-0.04		-0.23**	-0.22**		-0.03		-0.04	-0.03
		(0.10)		(0.11)	(0.11)		(0.05)		(0.08)	(0.08)
Log Male Wage (t)					0.23**					0.13**
					(0.11)					(0.08)
Log Female Wage (t)					-0.01					0.06**
					(0.08)					(0.04)
First-Stage:										
Fraction Female	0.51***	0.72***	0.58***	0.86***	0.85***	0.51***	0.72***	0.57***	0.84***	0.83***
	(0.11)	(0.12)	(0.13)	(0.14)	(0.14)	(0.11)	(0.12)	(0.13)	(0.14)	(0.14)
Reduced-Form:										
Fraction Female	-0.44	-0.52**	-1.21***	-1.26***	-1.11***	-0.47***	-0.60***	-0.63***	-0.85***	-0.72***
	(0.28)	(0.27)	(0.26)	(0.28)	(0.29)	(0.15)	(0.19)	(0.17)	(0.21)	(0.20)
Sample Size	1816	1816	1456	1456	1456	1816	1816	1466	1466	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$

TABLE V: TWO-STAGE LEAST-SQUARES ESTIMATES OF PERCENT FEMALE ON WAGE, 20-YEAR LAG

		Female Wage (t+20)			Male Wage (t+20)		
2SLS:		(1)	(2)	(3)	(4)	(6)	
	Fraction Female	-1.16**	-0.78**	-0.86***	-0.43	-0.46*	-0.45*
		(0.52)	(0.33)	(0.34)	(0.35)	(0.31)	(0.31)
	Labor Supply		-0.16**	-0.16**		0.01	0.02
			(0.09)	(0.09)		(0.09)	(0.08)
	Log Male Wage (t)			0.02			-0.09
				(0.08)			(0.09)
	Log Female Wage (t)			0.02			-0.09
				(0.08)			(0.09)
First-Stage:							
	Fraction Female	0.67***	0.91***	0.90***	0.61***	0.86***	0.83***
		(0.16)	(0.15)	(0.16)	(0.16)	(0.17)	(0.17)
Reduced-Form:							
	Fraction Female	-0.78***	-0.71***	-0.78***	-0.26	-0.40	-0.38
		(0.26)	(0.27)	(0.27)	(0.20)	(0.25)	(0.24)
	Sample Size	1097	1097	1097	1116	1116	1116

Note: Each column reports results from an estimate of equations (5), (6) and (7) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE VI: MALE AND FEMALE COLLEGE MAJOR CHOICE, 2009-2014

Major Field of Study	College Only %of Males	College Only %of Females	Master's or Higher %of Males	Master's or Higher %of Females
Medical and Health Services	2.4	11.0	3.3	8.9
Education	3.9	12.0	6.2	17.0
Social Work/Interdisciplinary/Psychology	5.8	12.0	6.8	16.0
Humanities/Communications/Law	15.0	20.0	12.0	16.0
Biological/Life Sciences	5.9	5.1	11.0	9.8
Business	27.0	23.0	16.0	12.0
Social Science	10.0	7.5	13.0	10.0
Construction/Manufacturing/Criminal Justice	4.7	2.1	1.6	1.1
Math/Computer Science	8.3	2.9	7.3	3.2
Physical Sciences/Engineering	17.0	4.7	23.0	7.0

*Note: Calculations from the 2009-2014 American Community Survey. Sample includes all respondents who include any major and any education. Includes only first major listed.

TABLE VII: ROBUSTNESS RESULTS FOR TWO-STAGE LEAST-SQUARES

	Log Female Wage				Log Male Wage			
	Major Instrument		1970 Base Year		Major Instrument		1970 Base Year	
	(t)	(t+10)	(t)	(t+10)	(t)	(t+10)	(t)	(t+10)
2SLS:								
Fraction Female	-2.13** (1.16)	-2.07* (1.53)	-0.62 (0.62)	-1.42** (0.62)	0.05 (0.61)	-0.16 (0.59)	-1.42*** (0.58)	-0.93** (0.44)
Labor Supply (1970)			-0.21** (0.12)	-0.44*** (0.15)			-0.22** (0.09)	-0.14* (0.09)
Log Male Wage		0.08 (0.17)		0.26*** (0.11)		0.17** (0.07)		0.12 (0.09)
Log Female Wage		0.08 (0.13)		-0.17** (0.07)		0.18*** (0.06)		0.04 (0.05)
First-Stage:								
Instrument	0.89*** (0.31)	0.81** (0.33)	0.61*** (0.14)	0.80*** (0.16)	0.89*** (0.31)	0.83** (0.33)	0.61*** (0.14)	0.79*** (0.16)
Sample Size	827	669	1348	1059	827	670	1348	1071

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. This table provides estimated effects using two alternative instrument definitions. "Major Instrument," the instrument in specifications 1, 2, 5, and 6, constructs the instrument by holding constant the fraction of workers from each college major entering an occupation, with a base year of 2010. Analysis using "Major Instrument" drops all occupations where fewer than 25% of workers held a bachelor's degree or higher in 2010. "1970 Base Year," the instrument in columns 3, 4, 7, and 8 uses 1970, rather than 1980 as a base year. These regressions drop the year 1960 from the analysis. The unit of observation is the occupation X year. Estimates control for age, education and base-year education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE VIII: EFFECT ON HOURS WORKED

	Log Female Wage		Log Male Wage	
2SLS:	(t)	(t+10)	(t)	(t+10)
Fraction Female	-0.95 (5.24)	-1.90 (4.03)	-4.00 (4.24)	-2.32 (3.29)
Labor Supply	-2.09** (1.21)	-1.55* (1.03)	-2.31*** (0.79)	-3.74*** (0.76)
Male Wage		-1.57* (0.98)		-0.78 (0.83)
Female Wage		-0.05 (0.87)		0.11 (0.31)
First-Stage:				
Fraction Female	0.72*** (0.12)	0.85*** (0.14)	0.72*** (0.12)	0.83*** (0.14)
Reduced-Form:				
Fraction Female	-0.68 (3.79)	-1.62 (3.45)	-2.89 (3.05)	-1.94 (2.76)
Sample Size	1816	1456	1816	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with average hours worked for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$

TABLE IX:

EFFECT ON ELASTICITY OF EARNINGS WITH RESPECT TO HOURS PER WEEK

	Elasticity of Earnings With Respect to Hours/Week		
2SLS:	(t)	(t+10)	(t+20)
Fraction Female	1.14**	1.03**	0.45
	(0.57)	(0.53)	(0.53)
Labor Supply	0.16*	0.11	0.04
	(0.12)	(0.13)	(0.12)
First-Stage:			
Instrument	0.78***	0.87***	0.86***
	(0.16)	(0.17)	(0.19)
Dependent Mean	0.26	0.36	0.42
Sample Size	1816	1478	1156

Note: Dependent Variable is the estimated elasticity of earnings with respect to hours worked for full-time workers in an occupation. This elasticity is given by $\beta j,t$ from equation (13). Regressions are weighted by the inverse standard error of $\beta j,t$. All specifications include occupation and year fixed-effects, time-varying controls and base-year education controls. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

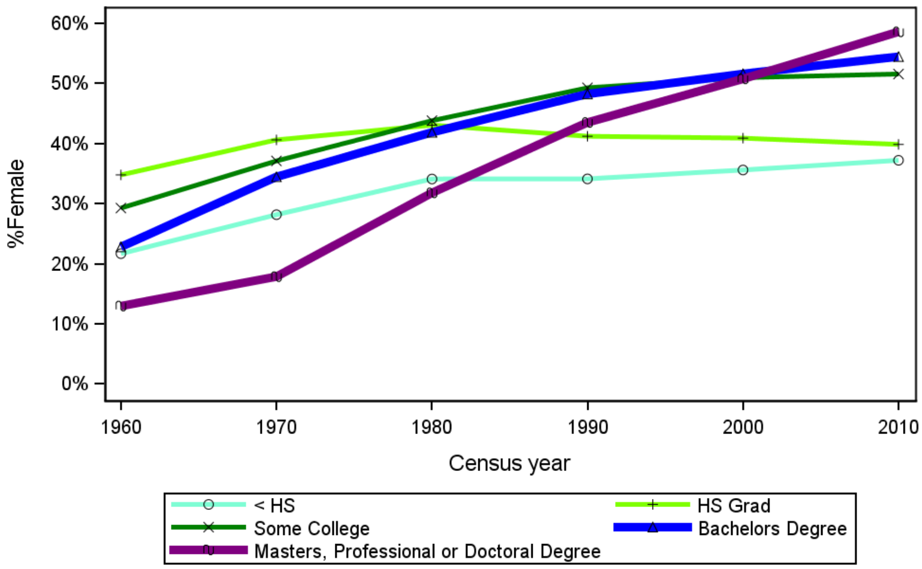
TABLE X: EFFECT ON PRESTIGE

	Log Mean Wage	Change in Prestige (1965-1989)	
2SLS:	(1)	(2)	(4)
Fraction Female	-33.04*	-0.87	-16.05
	(20.03)	(0.80)	(14.71)
Labor Supply	-0.24	0.04	-4.97**
	(2.64)	(0.10)	(2.94)
Base-Year Education		X	X
Time-Varying Controls			X
First-Stage:			
Instrument	0.43***	0.43***	0.80***
	(0.12)	(0.12)	(0.13)
Sample Size	302	302	302

Note: Dependent Variable is the estimated elasticity of earnings with respect to hours worked for full-time workers in an occupation. This elasticity is given by $\beta_{j,t}$ from equation (13). Regressions are weighted by the inverse standard error of $\beta_{j,t}$. All specifications include occupation and year fixed-effects, time-varying controls and base-year education controls. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

Figures

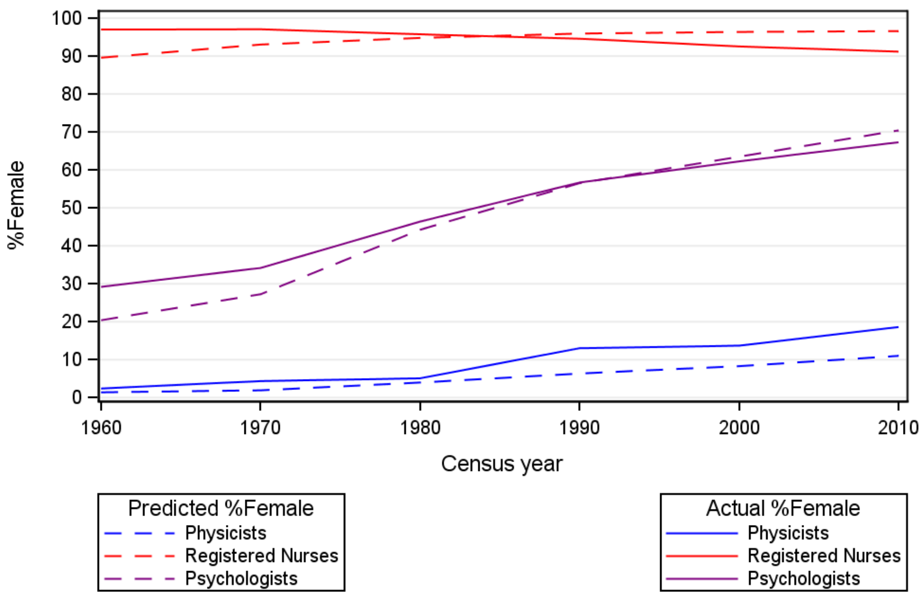
FIGURE I: GENDER COMPOSITION OF FULL-TIME WORKERS AGED 18-35 BY HIGHEST DEGREE



Note: This figure shows the percent of US workers of each education type aged 22-35 who are female in each census year from 1960-2010. Definitions of Education Type are available in the data appendix.

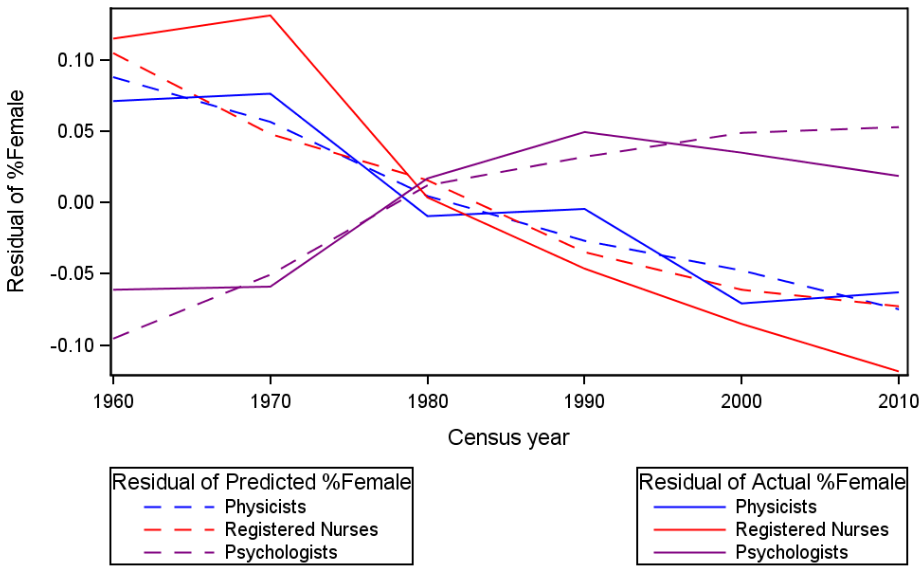
FIGURE II: PREDICTED AND ACTUAL PERCENT FEMALE

FOR THREE EXAMPLE OCCUPATIONS:



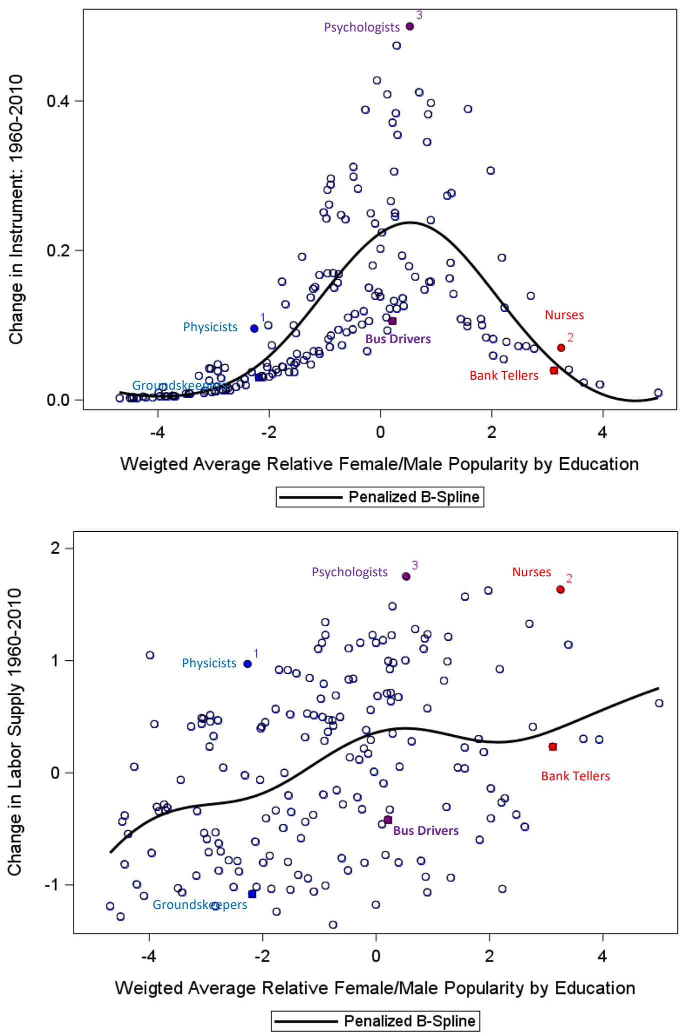
Note: This figure shows the percent female among workers in three occupations, as well as the instrumented percent female for young workers in those occupations

FIGURE III: RESIDUAL OF PREDICTED AND ACTUAL PERCENT FEMALE FOR THREE EXAMPLE OCCUPATIONS



Note: This figure shows the residual of percent female and instrumented percent female for workers in three occupations after controlling for occupation and year fixed effects and for controls described in the text.

FIGURE IV: RELATIONSHIP BETWEEN INSTRUMENT, LABOR SUPPLY AND RELATIVE FEMALE/MALE POPULARITY BY OCCUPATION



Notes: Weighted Average Relative Female/Male Popularity Represents the ratio of the percent of female workers of each education level in an occupation to the percent of male workers of each education level in an occupation, weighted by the percent of workers of the education level in the occupation. These graphs show the relationship between relative popularity and change from 1960-2010 in the instrument, induced total labor supply, induced male labor supply and induced female labor supply.

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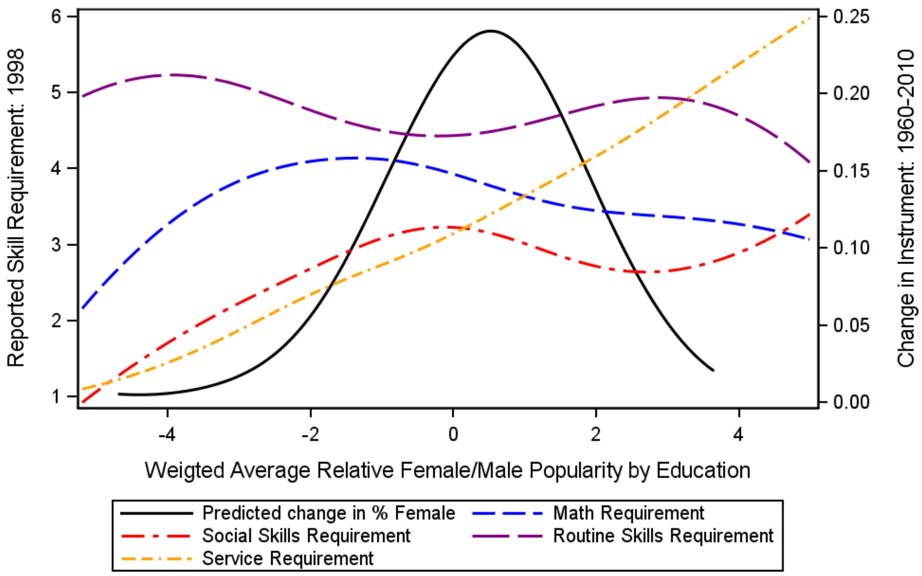
Physicists

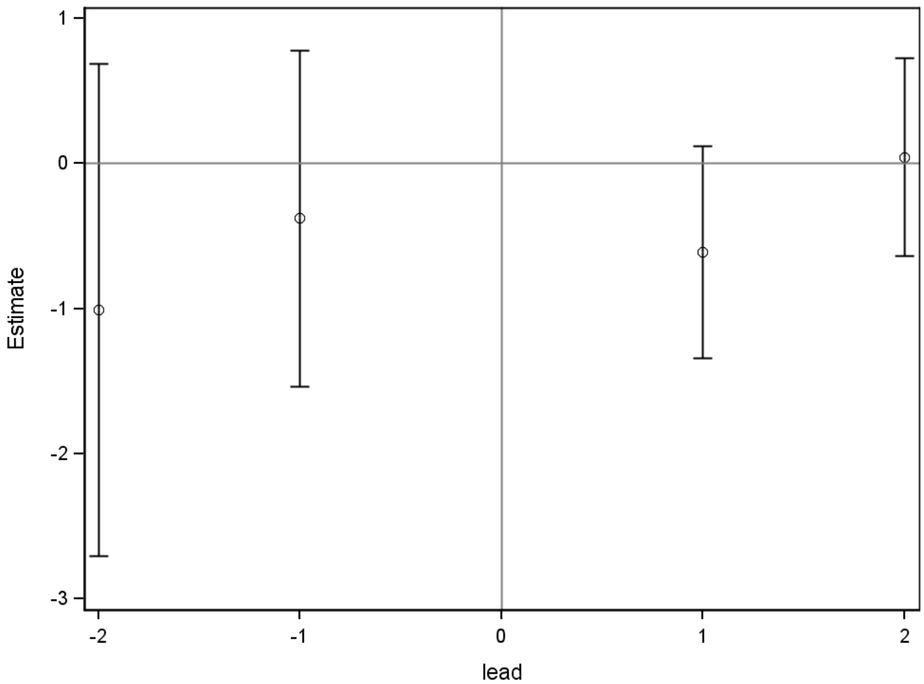
Bank Tellers

FIGURE V: RELATIONSHIP BETWEEN SKILL REQUIREMENTS AND WEIGHTED RELATIVE FEMALE/MALE POPULARITY BY OCCUPATION



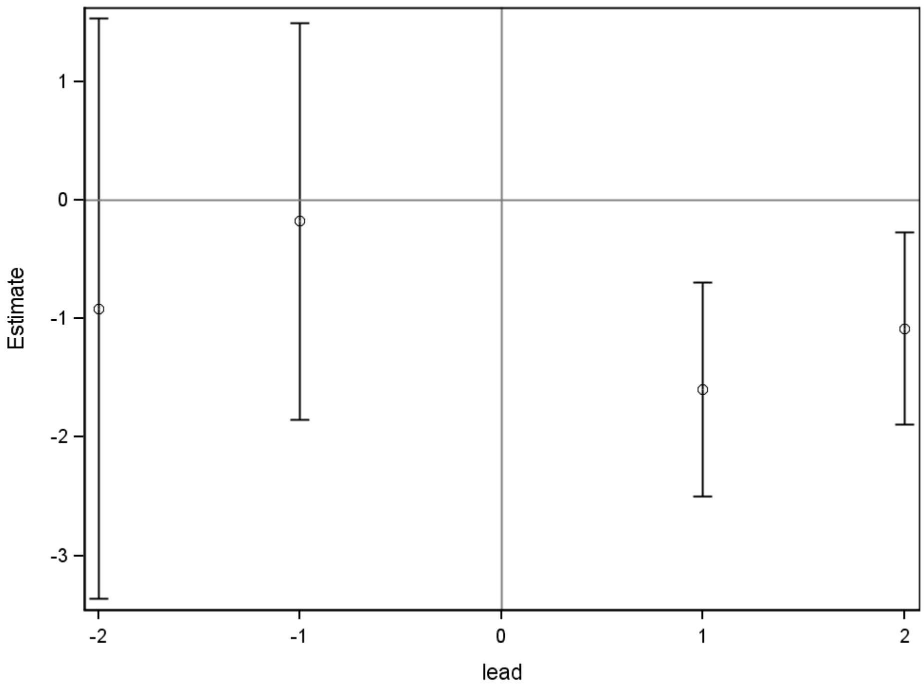
Note: Weighted Average Relative Female/Male Popularity by Education represents the ratio of the percent of female workers of each education level in an occupation to the percent of male workers of the education level in an occupation. This graph shows the nonparametric relationship between each skill and this average ratio, calculated using a penalized b-spline

FIGURE VI: EVENT STUDY WITH MALE WAGE AS DEPENDENT VARIABLE:



Notes: This figure presents estimated values and 95% confidence intervals of the coefficient β_1 in equation 8, estimated for female wages, with time lags of -20 years,
-10 years, 10 years, and 20 years. Each estimate reflects the effect of the instrument in year *t* on log mean male wage in year *t+k*, controlling for the instrument in year *t+k*. Standard errors are clustered at the occupation level.

FIGURE VII: EVENT STUDY WITH FEMALE WAGE AS DEPENDENT VARIABLE:



Notes: This figure presents estimated values and 95% confidence intervals of the coefficient β_1 in equation 8, estimated for female wages, with time lags of -20 years,
-10 years, 10 years, and 20 years. Each estimate reflects the effect of the instrument in year *t* on log mean female wage in year *t+k*, controlling for the instrument in year *t+k*. Standard errors are clustered at the occupation level.

SUPPLEMENTARY TABLES AND FIGURES

TABLE A.I: CROSS-SECTIONAL ASSOCIATION BETWEEN FRACTION FEMALE AND WAGE

	Log Female Wage			
	(1)	(2)	(3)	(4)
Fraction Female	-0.42*** (0.08)	-0.46*** (0.1)	-0.1 (0.07)	-0.15 (0.14)
Time-Varying Controls	X	X		
Weighted		X		X
Sample Size	294	294	321	321
	Log Male Wage			
	(1)	(2)	(3)	(4)
Fraction Female	-0.48*** (0.06)	-0.61*** (0.12)	0.06 (0.07)	-0.04 (0.14)
Time-Varying Controls	X	X		
Weighted		X		X
Sample Size	294	294	334	334

*Note: Each column reports the estimated effect of fraction female on log wage for workers in the 2010 American Community Survey. Time-varying controls are age and education, as described in the text. Weighted regressions are weighted by the number of total workers in the occupation in 2010. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.II: RESULTS WEIGHTED BY SQUARE ROOT OF NUMBER OF RESPONDENTS

	Females				Males			
	Log Wage (t)	Log Wage (t+10)			Log Wage (t)	Log Wage (t+10)		
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fraction Female	-0.31	-0.85**	-0.86**	-1.40**	-0.25	-0.42*	-0.47*	-0.95**
	(0.38)	(0.41)	(0.40)	(0.70)	(0.26)	(0.26)	(0.28)	(0.47)
Labor Supply	-0.06	-0.15*	-0.16*		-0.03	-0.12**	-0.13**	
	(0.10)	(0.10)	(0.10)		(0.05)	(0.07)	(0.07)	
Male Wage		0.20**				0.15**		
		(0.10)				(0.07)		
Female Wage		-0.00				0.05*		
		(0.07)				(0.03)		
First-Stage:								
Fraction Female	0.71***	0.78***	0.79***	0.49***	0.71***	0.77***	0.78***	0.45**
	(0.19)	(0.20)	(0.19)	(0.19)	(0.19)	(0.20)	(0.20)	(0.19)
Sample Size	1816	1456	1456	1456	1816	1466	1466	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Each observation is weighted by the square root of the number of respondents with the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.III: OCCUPATIONS WITH HIGHEST PREDICTED CHANGE IN PERCENT FEMALE FROM 1960-2010

Occupation	Δ Resid. Instrument	Δ Resid. %Female	% Female (1960)	Δ % Female, 1960-2010	%College 1980	Weighted Average 1980 Gender Ratio
Social scientists, n.e.c.	0.147	1%	40%	15%	85%	-0.07
Psychologists	0.146	8%	29%	38%	94%	0.52
Subject instructors (HS/college)	0.134	-6%	23%	22%	92%	0.28
Therapists, n.e.c.	0.130	-3%	59%	19%	70%	0.85
Managers in education and related fields	0.124	22%	22%	42%	74%	0.12
Art/entertainment performers and related	0.119	4%	21%	25%	35%	-0.18
Social workers	0.115	0%	61%	16%	69%	0.90
Editors and reporters	0.107	-1%	35%	17%	72%	0.30
Writers and authors	0.105	17%	17%	41%	76%	0.27
Cooks, variously defined	0.100	-21%	62%	-14%	4%	0.78
Shoemaking machine operators	0.094	67%	3%	61%	1%	1.31
Misc textile machine operators	0.094	-34%	78%	-32%	2%	-0.01
Economists, market researchers, and survey researchers	0.090	-1%	13%	20%	79%	-0.49

Occupation	Δ Resid. Instrument	Δ Resid. %Female	% Female (1960)	Δ % Female, 1960-2010	%College 1980	Weighted Average 1980 Gender Ratio
Insurance adjusters, examiners, and investigators	0.090	41%	13%	62%	33%	0.90
Managers of properties and real estate	0.088	0%	29%	23%	33%	-0.14
Teachers , n.e.c.	0.087	-10%	55%	3%	56%	0.83
Designers	0.087	19%	15%	37%	42%	0.18
Art makers: painters, sculptors, craft-artists, and print-makers	0.087	-12%	31%	7%	45%	0.24
Other financial specialists	0.086	8%	24%	33%	48%	0.26
Real estate sales occupations	0.084	8%	22%	29%	38%	0.25

TABLE A.IV: OCCUPATIONS WITH LOWEST PREDICTED CHANGE IN PERCENT FEMALE FROM 1960-2010

Occupation	Δ Resid. Instrument	Δ Resid. %Female	% Female (1960)	Δ % Female, 1960-2010	%College 1980	Weighted Average 1980 Gender Ratio
Registered nurses	-0.175	-23%	97%	-6%	42%	3.25
Mechanical engineers	-0.164	-11%	0%	8%	66%	-3.91
Chemical engineers	-0.161	-10%	1%	14%	87%	-3.08
Civil engineers	-0.159	-9%	1%	13%	74%	-3.27
Physicists and astronomers	-0.158	-13%	2%	16%	87%	-2.27
Aerospace engineer	-0.156	-12%	1%	11%	77%	-3.05
Electrical engineer	-0.142	-11%	1%	10%	70%	-2.80
Dentists	-0.131	-7%	5%	22%	98%	-1.56
Dietitians and nutritionists	-0.126	-20%	91%	0%	61%	2.70
Not-elsewhere-classified engineers	-0.125	-12%	0%	12%	71%	-2.93
Insulation workers	-0.110	6%	2%	10%	4%	-2.70
Licensed practical nurses	-0.109	-13%	95%	-2%	7%	3.38
Optometrists	-0.092	10%	5%	37%	96%	-1.72
Industrial engineers	-0.086	-6%	2%	16%	56%	-1.96
Architects	-0.073	0%	2%	24%	83%	-1.78
Foresters and conservation scientists	-0.072	-2%	2%	18%	63%	-2.02
Typists	-0.070	-24%	95%	-5%	7%	3.65
Excavating and loading machine operators	-0.066	2%	0%	1%	1%	-4.10

Occupation	Δ Resid. Instrument	Δ Resid. %Female	% Female (1960)	Δ % Female, 1960-2010	%College 1980	Weighted Average 1980 Gender Ratio
Power plant operators	-0.062	-10%	5%	2%	8%	-2.94
Misc material moving occupations	-0.058	-6%	1%	1%	4%	-1.55

TABLE A.V: MAIN RESULTS WITH SKILL X YEAR CONTROLS

	Females				Males			
	Wage (t)		Wage (t+10)		Wage (t)		Wage (t+10)	
2SLS:	(1)	(2)	(3)	-4	(5)	(6)	(7)	(8)
Fraction Female	-0.74** (0.37)	-0.56* (0.38)	-1.30*** (0.39)	-1.30*** (0.39)	-0.78*** (0.26)	-0.74** (0.37)	-0.78*** (0.25)	-0.68*** (0.25)
Labor Supply	-0.04 (0.10)	-0.01 (0.10)	-0.22** (0.11)	-0.19** (0.11)	-0.01 (0.06)	-0.04 (0.10)	-0.03 (0.08)	0.00 (0.09)
Male Wage			0.21** (0.12)	0.24** (0.12)			0.15** (0.09)	0.14** (0.08)
Female Wage			-0.02 (0.08)	-0.02 (0.08)			-0.03 (0.08)	0.06** (0.03)
Skill X Year		X		X		X		X
First-Stage:								
Inst.	0.75*** (0.13)	0.77*** (0.14)	0.91*** (0.15)	0.95*** (0.16)	0.75*** (0.13)	0.77*** (0.14)	0.89*** (0.15)	0.94*** (0.17)
Sample Size	1575	1575	1276	1276	1575	1575	1283	1283

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.VI: MAIN RESULTS CONTROLLING FOR FUTURE VALUE OF INSTRUMENT

	Log Female Wage		Log Male Wage	
2SLS:	(1)	(2)	(3)	(4)
Fraction Female	-0.57 (0.57)	-0.71 (0.65)	-0.74** (0.37)	-0.70* (0.46)
Instrument (t+10)	-0.19 (0.56)	0.67 (1.65)	-0.05 (0.39)	1.03 (0.97)
Instrument (t+20)		-1.52 (2.30)		-1.57 (1.36)
Labor supply	-0.01 (0.16)	-0.01 (0.17)	-0.01 (0.09)	-0.02 (0.09)
First-Stage:				
Fraction Female	1.11*** (0.29)	1.19*** (0.28)	1.04*** (0.28)	1.06*** (0.28)
Reduced-Form:				
Fraction Female	-0.63 (0.65)	-0.85 (0.83)	-0.78** (0.36)	-0.74 (0.46)
Sample Size	1456	1097	1466	1116

Note: Note: Each column reports results from an estimate of equations (5) (6) and (7) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. I add an estimate of the instrument at t+10 and t+20. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.VII: MAIN RESULTS ESTIMATED FOR AGGREGATED OCCUPATIONS

	Females			Males		
	Log Wage (t)	Log Wage (t+10)		Log Wage (t)	Log Wage (t+10)	
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)
Fraction Female	-0.80*	-1.39***	-1.42***	-0.58	-0.52*	-0.45*
	(0.50)	(0.53)	(0.55)	(0.47)	(0.35)	(0.30)
	-0.14	-0.34**	-0.34**	-0.00	-0.15*	-0.14**
ls	(0.14)	(0.16)	(0.16)	(0.12)	(0.09)	(0.08)
Male Wage (t)			-0.18			0.04
			(0.16)			(0.18)
Female Wage (t)			0.11			0.11
			(0.12)			(0.11)
First-Stage:						
Fraction Female	0.80***	0.91***	0.91***	0.80***	0.94***	0.93***
	(0.27)	(0.26)	(0.26)	(0.27)	(0.26)	(0.25)
Reduced-Form:						
Fraction Female	-0.64**	-1.27***	-1.28***	-0.46	-0.49	-0.42
	(0.31)	(0.35)	(0.37)	(0.33)	(0.30)	(0.27)
Sample Size	458	379	379	458	381	381

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is Aggregated occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.VIII: CORRELATIONS BETWEEN RESIDUAL OF PRIMARY AND ROBUSTNESS INSTRUMENTS

	1980 Base Year	1970 Base Year	Major Instrument
1980 Base Year	1.00 (3634)		
1970 Base Year	0.85 (3114)	1.00 (3114)	
Major Instrument	0.35 (3126)	0.32 (2708)	1.00 (3126)

Note: The following table shows the correlations between the instrument defined using a 1980 base year, the instrument defined using a 1970 base year, and the instrument defined using a 2010 base year and including college major, after controlling for occupation and year fixed effects, current-period education and age, time-varying base-year education and induced labor supply. Numbers in parenthesis are sample sizes

TABLE A.IX: PRIMARY RESULTS ON LOG MEAN WAGE OF COLLEGE GRADUATES

	Females				Males			
	Log Wage (t)	Log Wage (t+10)			Log Wage (t)	Log Wage (t+10)		
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fraction Female	-0.75 (0.5)	-0.84* (0.43)	-0.74* (0.44)	-1.92*** (0.66)	-1.14*** (0.39)	-0.89** (0.35)	-0.71** (0.35)	-1.00** (0.5)
Labor Supply	-0.25** (0.1)	-0.57*** (0.11)	-0.55*** (0.11)		-0.15* (0.08)	-0.22** (0.09)	-0.19** (0.09)	
Female Labor Supply				-1.18*** (0.24)				-0.78*** (0.19)
Male Labor Supply				0.24 (0.26)				0.21 (0.19)
Log Male Wage (t)		-0.14*** (0.03)				-0.15*** (0.03)		
First-Stage:								
Fraction Female	0.71*** (0.07)	0.85*** (0.09)	0.84*** (0.09)	0.54*** (0.08)	0.71*** (0.07)	0.86*** (0.09)	0.84*** (0.09)	0.52*** (0.08)
Reduced-Form:								
Fraction Female	-0.53 (0.35)	-0.71** (0.36)	-0.63* (0.36)	-1.03*** (0.31)	-0.81*** (0.26)	-0.76*** (0.28)	-0.59** (0.28)	-0.52** (0.24)
Sample Size	1495	1155	1155	1155	1495	1155	1155	1155

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 with a bachelor's degree or above as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.X: PRIMARY RESULTS ON WORKERS AGE 22-35,

NOT ENROLLED IN SCHOOL

	Females			Males		
	Log Wage (t)	Log Wage (t+10)		Log Wage (t)	Log Wage (t+10)	
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)
Fraction Female	-0.98*** (0.35)	-1.12*** (0.37)	-1.34*** (0.39)	-0.57** (0.32)	-0.28 (0.23)	-0.50** (0.26)
Labor Supply	-0.25*** (0.07)	-0.31*** (0.09)	-0.32*** (0.10)	-0.15*** (0.05)	-0.16*** (0.04)	-0.16*** (0.05)
Log Male Wage (t)		0.32*** (0.08)			0.24*** (0.05)	
Log Female Wage (t)		-0.02 (0.06)			0.05** (0.02)	
First-Stage:						
Fraction Female	0.72*** (0.12)	0.85*** (0.14)	0.86*** (0.14)	0.72*** (0.12)	0.83*** (0.14)	0.84*** (0.14)
Reduced-Form:						
Fraction Female	-0.71*** (0.22)	-0.95*** (0.25)	-1.15*** (0.26)	-0.41* (0.22)	-0.23 (0.19)	-0.42** (0.20)
Sample Size	1816	1456	1456	1816	1466	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers who are not in school and are age 22-35 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.XI: MAIN RESULTS WITHOUT LABOR SUPPLY CONTROL,

WITHOUT BASE-YEAR EDUCATION COMPOSITION CONTROL

	Females						Males					
	Log Wage (t)			Log Wage (t+10)			Log Wage (t)			Log Wage (t+10)		
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Fraction	-0.72**	-0.86*	-0.34	-1.30***	-1.94***	-1.13**	-0.82***	-0.92***	-0.42*	-0.87***	-0.97***	-0.42*
Female	(0.39)	(0.59)	(0.45)	(0.42)	(0.72)	(0.53)	(0.28)	(0.34)	(0.27)	(0.27)	(0.37)	(0.31)
Labor	-0.04			-0.22**			-0.03			-0.03		
Supply	(0.10)			(0.11)			(0.05)			(0.08)		
Log Male				0.23**	0.21*	0.23**				0.13**	0.13*	0.13**
Wage (t)				(0.11)	(0.13)	(0.10)				(0.08)	(0.08)	(0.07)
Log												
Female				-0.01	-0.02	0.02				0.06**	0.06*	0.08***
Wage (t)				(0.08)	(0.08)	(0.07)				(0.04)	(0.04)	(0.04)
Base-year												
Education	X	X		X	X		X	x		X	X	
First-Stage:												
Fraction	0.72***	0.51***	0.54***	0.85***	0.57***	0.57***	0.72***	0.51***	0.54***	0.83***	0.56***	0.56***
Female	(0.12)	(0.11)	(0.10)	(0.14)	(0.13)	(0.13)	(0.12)	(0.11)	(0.10)	(0.14)	(0.13)	(0.13)
Reduced-Form:												
Fraction	-0.52**	-0.44	-0.18	-1.11***	-1.10***	-0.65***	-0.60***	-0.47***	-0.23*	-0.72***	-0.54***	-0.23
Female	(0.27)	(0.28)	(0.23)	(0.29)	(0.27)	(0.24)	(0.19)	(0.15)	(0.14)	(0.20)	(0.16)	(0.16)
Sample												
Size	1816	1816	1816	1456	1456	1456	1816	1816	1816	1466	1466	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Base-Year Controls are the share of the workforce in each education category in 1980, interacted with year. Standard errors are in parenthesis and are clustered at the occupation level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$

TABLE A.XII: MAIN RESULTS ESTIMATED WITH CONTROLS FOR EMPLOYMENT

	Females				Males			
	Log Wage (t)	Log Wage (t+10)			Log Wage (t)	Log Wage (t+10)		
2SLS:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fraction Female	-0.73**	-1.30***	-1.47***	-1.88***	-0.78***	-0.81***	-0.95***	-0.99***
	(0.39)	(0.43)	(0.43)	(0.62)	(0.27)	(0.25)	(0.28)	(0.33)
Labor Supply	-0.01	-0.23**	-0.23**		-0.04	-0.04	-0.05	
	(0.09)	(0.11)	(0.11)		(0.06)	(0.08)	(0.08)	
Log Employment (t)	0.02				-0.02*			
	(0.02)				(0.01)			
Log Employment (t+10)		-0.00	-0.00	0.01		-0.03***	-0.04***	-0.04**
		(0.02)	(0.02)	(0.02)		(0.01)	(0.02)	(0.02)
Male Wage		0.23**				0.13**		
		(0.11)				(0.08)		
Female Wage		-0.01				0.06**		
		(0.08)				(0.04)		
First-Stage:								
Fraction Female	0.72***	0.84***	0.84***	0.61***	0.76***	0.87***	0.88***	0.60***
	(0.11)	(0.14)	(0.13)	(0.13)	(0.11)	(0.14)	(0.13)	(0.12)
Reduced-Form:								
Fraction Female	-0.52**	-1.09***	-1.24***	-1.15***	-0.60***	-0.71***	-0.83***	-0.60***
	(0.27)	(0.29)	(0.28)	(0.27)	(0.19)	(0.20)	(0.21)	(0.16)
Sample Size	1816	1456	1456	1456	1816	1466	1466	1466

Note: Each column reports results from an estimate of equations (14) and (15) in the paper, with Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation X year. Estimates control for age, education and 1980 education of workers, as described in the text. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.XIII: 2SLS ESTIMATE OF THE LONG-TERM EFFECT OF GENDER COMPOSITION ON WAGE

First-Stage:	Δ percent female	Δ percent female
Δ Instrumented percent female	0.60** (0.24)	0.60** (0.24)
Δ Induced LSF	-0.10 * (0.06)	-0.10 * (0.06)
Reduced-Form:	Δ Female Wage	Δ Female Wage
Δ Instrumented percent female	-0.55 (0.46)	-0.21 (0.27)
Δ Induced LSF	0.17 (0.14)	0.09 (0.08)
2SLS:	Δ Female Wage	Δ Male Wage
Δ Instrumented percent female	**-0.91** **(0.87)**	**-0.36** **(0.45)**
Δ Induced LSF	0.08 (0.13)	0.05 (0.07)
Sample Size:	160	160

*Note: Each column reports results from an estimate of equations (16) and (17) in the paper, with the change from 1960-2010 in Log Mean Wage for workers over the age of 45 as the dependent variable, estimated for males and females. The unit of observation is the occupation. Controls are described in the paper, and are measured as the difference between 1960 and 2010 values. Standard errors are in parenthesis and are clustered at the occupation level. *** p<0.01, ** p<0.05, * p<0.10

TABLE A.XIV: MAJORS IN THE AMERICAN COMMUNITY SURVEY

major	**Disaggregated majors included**	**#with major (%with major) among bachelors only**	**#with major (%with major) among post-bachelors degrees**
1: Math/Computer Science	20: Communication Technologies, 21: Computer and Information Sciences, 37: Mathematics and Statistics	70593 (5.3%)	36584 (4.9%)
2: Humanities/ Communications/ Law	19: Communications, 26: Linguistics and Foreign Languages, 32: Law, 33: English Language, Literature, and Composition, 34: Liberal Arts and Humanities, 35: Library Science, 48: Philosophy and Religious Studies, 49: Theology and Religious Vocations, 60: Fine Arts	236101 (17.9%)	103634 (14%)
3: Biological/ Life Sciences	11: Agriculture, 13: Environment and Natural Resources, 36: Biology and Life Sciences	72141 (5.5%)	75984 (10.3%)
4: Medical and Health Services	61: Medical and Health Sciences and Services	91937 (7%)	47942 (6.5%)
5: Social Work/ Interdisciplinary/ Psychology	15: Area, Ethnic, and Civilization Studies, 22: Cosmetology Services and Culinary Arts, 29: Family and Consumer Sciences, 40: Interdisciplinary and Multi-Disciplinary Studies (General), 41: Physical Fitness, Parks, Recreation, and Leisure, 52: Psychology, 54: Public Affairs, Policy, and Social Work	122429 (9.3%)	88923 (12%)
6: Physical Sciences/ Engineering	14: Architecture, 24: Engineering, 25: Engineering Technologies, 50: Physical	135441 (10.3%)	102661 (13.9%)

major	Disaggregated majors included	#with major (%with major) among bachelors only	#with major (%with major) among post-bachelors degrees
	Sciences, 51: Nuclear, Industrial Radiology, and Biological Technologies		
7: Construction/ Manufacturing/ Criminal Justice	38: Military Technologies, 53: Criminal Justice and Fire Protection, 56: Construction Services, 57: Electrical and Mechanic Repairs and Technologies, 58: Precision Production and Industrial Arts, 59: Transportation Sciences and Technologies	43179 (3.3%)	9792 (1.3%)
8: Education	23: Education Administration and Teaching	111569 (8.4%)	90436 (12.2%)
9: Social Science	55: Social Sciences, 64: History	114824 (8.7%)	86304 (11.7%)
10: Business	62: Business	322958 (24.4%)	97929 (13.2%)

FIGURE A.I: RELATIONSHIP BETWEEN WEIGHTED FEMALE/MALE POPULARITY RATIO AND SKILLS