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Does bad pay cause occupations to feminize, Does feminization reduce pay, and How can we tell with longitudinal data?

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Abstract

Predominantly female occupations pay less than "male" occupations, even after adjusting for skill demands. The devaluation perspective says that this is because gender bias influences employers' decisions about the relative pay of occupations. The longitudinal implication of the devaluation perspective is that when occupations feminize, their relative pay should go down. In the queueing or relative-attractiveness view, occupations' reward levels affect their sex composition, with less attractive occupations going to women because employers prefer men and can get them in occupations that pay well. To test these views about change over time, we use a fixed-effects model with lagged independent variables and data from the 1983–2001 Current Population Surveys. There is only slight evidence that the feminization of occupations lowers their wages, and no evidence that a fall in occupations' relative wages leads to feminization. But percent female is associated with lower wages in every year. We speculate that in the early history of new occupations and organizations, there was a causal relationship between gender composition and wages, and it has been frozen in by institutional inertia in relative wage structures.

Keywords: Gender inequality; Occupations; Discrimination; Earnings; Gender; Longitudinal analysis

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1. Introduction

In the United States, some of the pay gap between men and women comes from women's concentration in lower paying occupations (Petersen and Morgan, 1995; England, 1992), despite the fact that women's occupations require about as much education and skill as men's, on average. This correlation between occupations' sex composition and their average pay is well known, and holds up in the presence of numerous controls in cross-sectional data. There are two major sociological views of the causal dynamics involved, and they have different implications for how the sex composition and pay of occupations should co-vary over time. Past research has inadequately exploited longitudinal data to examine this relationship over time.

In the "devaluation" view, associated with policy proposals for "comparable worth," predominantly female occupations are paid less *because* women fill the occupations (England, 1992; Steinberg, 2001; Sorensen, 1994). In this view, if an occupation is filled mostly by women, employers see it as less valuable, less demanding, or less pay-worthy. Somehow, the low status of women "rubs off" on employers' evaluation of the occupation, and they set a lower pay level for both men and women in the occupation than they would have if the identical occupation were done mostly by men. If this bias in wage setting exists, then we would expect that if the sex composition of an occupation changes, its pay would change. The hypothesis would be appropriately tested with controls for skill demands of the occupations. But little of the analysis offered in support of the devaluation claim actually examines how *changes* in occupations' sex composition relate to *changes* in their pay.

Another view sees the causal arrow the opposite way, implying that bad pay (or other undesirable nonpecuniary characteristics) causes occupations to feminize. This view is associated with Reskin and Roos (1990) and with Catanzarite, Strober, and Arnold (Strober, 1984; Strober and Arnold, 1987; Strober and Catanzarite, 1994). In this view, employers' preferences for men, combined with men and women's preferences for better paying occupations, leads desirable occupations to be filled disproportionately by men. Given workers' preferences for good pay, when hiring for high paying occupations, employers will often be able to get men, but when hiring in low paying occupations, they will often have to settle for women even if they prefer men, since men will gravitate first to the high paying occupations. In such a process, even though women also prefer high paying occupations, they will only be able to get the occupations men do not want. Like devaluation, this hypothesis is appropriately tested controlling for educational and skill requirements. It is when occupations pay badly *relative to* their educational requirements that employers will be less able to get men, and the occupations are more likely to end up filled with women. Reskin and Roos' refer to this as the "queueing" view. Strober and Catanzarite (1994) have referred to it as the "relative attractiveness" theory of segregation; the more attractive an occupation is, the more likely it is to come to be filled by men. If this view is correct, longitudinal analysis should reveal that occupations' pay at one time affects their sex composition at a later time. As with the previous hypothesis, little of the analysis offered in support of this claim has actually examined how *changes* in occupations' sex composition relate to *changes* in their pay.

These two views both posit sex discrimination, but of different types. In the queueing or "relative attractiveness" view, it is hiring or placement discrimination against women. Such discrimination has been illegal since the Civil Rights Act of 1964 but undoubtedly still exists to some degree. In the devaluation view, gender bias affects which occupations get assigned higher wages, but this form of bias does not violate current law in the US.[1] With few exceptions, violations of the principle of ''comparable worth''—that occupations requiring the same amount of skill and having equally onerous working conditions must pay equivalently when in the same establishment—have not been found by US courts to violate antidiscrimination laws (Nelson and Bridges, 1999; England, 1992, Chapter 5). A third type of discrimination (which is illegal), paying women less than men in the same job, will not be our focus here.

The two theoretical views of the link between an occupation's sex composition and its wages make different predictions about how the two factors would covary over time. In the devaluation view, earlier levels of occupational sex composition should affect later wages; we would expect that changes in the sex composition would produce changes in pay. In the queueing or relative attractiveness view, earlier wages should affect later sex composition; we would expect that changes in wages would produce changes in sex composition. It is also possible that both devaluation and queueing processes could be going on simultaneously; the causal arrow may run both ways. Indeed, authors advocating the queueing view recognize the possibility of devaluation occurring during and following occupational feminization, and authors advocating the devaluation view recognize that there may be hiring or placement discrimination. In this case, changes in either factor will affect the other; it is the magnitude of each effect that is of interest.

Against both the devaluation and queueing view, two other theoretical positions lead to predictions of no causal effect between occupations' sex composition and their pay in either direction over time. The first of these is favored by many economists and by sociologist Tony Tam (1997). In this view, women choose less demanding occupations than men because they prioritize motherhood more and money less than men, and occupations into which women cluster pay less because they are less onerous or more ''mother-friendly.'' The claim is that women's occupations have more flexible hours, or allow a parent to use the phone to check in with children, or provide child care. Or, as Tam (1997) emphasizes, they may require less specialized training (occupation- or firm-specific) of the type that mothers who plan some time out of employment for childrearing do not find a worthwhile investment. These gender-specific hypotheses are consistent with a larger theory in neoclassical labor economics known as ''equalizing differentials'' which subsumes both human capital theory and the theory of compensating differentials (Rosen, 1986). The idea is that employers have to pay more to fill occupations that have nonpecuniary characteristics that workers do not like or that require them to bear costs (e.g. for training) to enter. Putting it statistically, this suggests that if we include the right control variables for

[1] Some economists agree that female occupations pay less than equally skilled male occupations, but disagree that employers' discriminatory assignment of lower wages because an occupation is female is the culprit. They believe that the problem is that female occupations are ''crowded'' and this excess of supply drives wages down. In Bergmann's (1971) formulation of the crowding hypothesis, employers will not hire women in male occupations, as in the Reskin and Roos (1990) or Strober and Catanzarite (1994) view. In this version of the crowding view, hiring discrimination in male jobs indirectly lowers wages in female jobs because it leads to overcrowding of the female jobs (by women who cannot get into the male jobs). Other economists agree with Bergmann (1971) that female jobs pay less because of crowding, but see the causes of crowding not in hiring discrimination but rather on the supply side. For example, if women prefer female gender-typed jobs, and there is an exogenous supply-side increase in female labor supply, this will lead to increased supply to female jobs that will lower their wages, even in the absence of any kind of discrimination. (Macpherson and Hirsch (1995) mention this as one possibility.)

occupational characteristics and human capital requirements, the relationship between the percentage female employees in an occupation and its pay will disappear, in either a cross-sectional or longitudinal analysis.

For entirely different reasons, institutional theories in sociology and economics cast doubt on the idea that changes in either sex composition or wages would flexibly adjust to each other. Institutionalist economists' studies of wage systems emphasize that hierarchies of the *relative* pay levels of occupations are surprisingly rigid. This is an example of the force of institutional and organizational inertia and path-dependence that is emphasized by scholars of organizations such as Stinchcombe (1965), and is agreed upon by both institutionalists and population ecologists among contemporary schools of organizational thought. Sociologists in the population–ecology camp have emphasized the long-term effects of "birth marks" on firms (Baron et al., 2002). Harmonizing with sociological institutionalism, institutionalist economists argue that relative wages of occupations within an organization often change little (Levine et al., 2002). In this view, relative wages of occupations will be very "sticky"—unlikely to move and then affect sex composition, and unlikely to move in response to sex composition.

In this paper, we use longitudinal data on occupations in the U.S. over the 19 year period from 1982 to 2000 to assess whether change in sex composition leads to change in pay or vice versa. Our main contribution is to shed light on the substantive question with a systematic longitudinal analysis, given the dearth of such analyses in past literature. However, in order to do this, we must grapple with the question of what statistical models are best for isolating causal effects in panel data, in the presence of substantial possible omitted variable bias. We will argue for using a fixed-effects model with lagged independent variables. We see our approach as an advance over the cross-lagged panel approach, sometimes called the "lagged-Y regressor model," which has typically been used in earlier quantitative studies on this topic. The fixed-effects approach provides superior protection against omitted variable bias by estimating the coefficients as if dummies for each occupation are controlled, thus controlling for all unmeasured, stable characteristics of occupations. We lag the independent variable a number of years behind the dependent variable. By running alternative models that reverse whether occupations' pay or sex composition is the dependent variable, with fixed-effects in the model in both instances, we can isolate causal effects in both directions with substantial protection from omitted variable bias. Our approach is also an advance over past research on this question in that we pool multiple years of data. Although the lagged-Y regressor approach can be done with pooled data, this has not been done in past studies using the lagged-Y regressor approach, all of which use just pairs of years.[2]

[2] Some longitudinal analyses have used person-years as observations and used *individual* fixed-effects models to examine how individuals' wages change when they move across occupations differing in sex composition. These analyses have shown some relationship between occupations' sex composition and wage even after adjusting for unmeasured individual characteristics via fixed effects (Kilbourne et al., 1994; Macpherson and Hirsch, 1995). Individual fixed-effects models have also cast doubt on the claim that female occupations are more mother friendly; Budig and England (2001) find no evidence that the wage penalty for motherhood is explained by mothers disproportionately working in female occupations. While these analyses share the advantage of the fixed-effects models proposed here for removing omitted variable bias, it is a different bias that they remove; they control for unmeasured, unchanging properties of the *individuals* who select into (or are selected by employers into) more male and female occupations. They do not give us purchase on whether unmeasured characteristics of occupations account for the relationship between their wages and their percent female.

2. Past research and theory on the causal relationship between occupational sex composition and wages

The queueing and relative attractiveness views were developed in large part based on longitudinal case studies. Strober and Arnold (1987) studied post-war change in bank telling, while Reskin and Roos (1990) did case studies of a number of occupations that increased their percent female by at least ten percentage points between 1970 and 1980.[3] In these case studies, there was evidence that occupations feminized when they were losing relative status or income potential.

Much of the support for devaluation or compensating differentials comes from quantitative cross-sectional analyses. Some cross-sectional regressions find that, net of skill demands, occupational percent female is associated with lower pay, consistent with devaluation (Sorensen, 1994; England, 1992; England et al., 2000; Jacobs and Steinberg, 1990a,b). Other cross-sectional regressions find no effect of sex composition on wages (Filer, 1985; Tam, 1997[4]).

We focus our review here on studies that are, like ours, quantitative and longitudinal, but that used a methodological approach that we will argue to be less appropriate than what we propose.[5] Five studies have used longitudinal data on a range of occupations or jobs in the US to investigate the causal order between the sex composition of occupations and their wages. A sixth similar analysis uses Israeli data. All these studies have used a pair of years and some version of a cross-lagged panel (lagged-Y regressor) model.

In a study of college administrators, Pfeffer and Davis-Blake (1987) concluded that there was evidence for causality in both directions. They used data for the two academic years ending in 1979 and 1984 from the College and University Personnel Association's Annual Administration Compensation Surveys on administrators' salaries. Consistent with devaluation, salaries decreased in response to increases in the proportion of women administrators in the institution (at least up to a relatively high point of percent female).

[3] While Reskin and Roos (1990) emphasized diminishing rewards as the major reason that occupations feminize, this is not the only factor they discussed. They also discussed examples of highly skilled, quickly growing occupations in which employers turned to women because of a shortage of men with the appropriate credentials.

[4] In a comment responding to Tam, England et al. (2000) replicated his models, adding the Dictionary of Occupational Titles' measure for the general educational requirement of occupations, and found that the negative effect of percent female reappeared. (See Tam, 2000 for a response.) In another article responding to Tam (1997), Tomaskovic-Devey and Skaggs (2002) used cross-sectional data from North Carolina and a second-stage-least squares approach. They argue that percent female affects wage, but by reducing firm-specific training in women's occupations, more than directly. Their analysis is distinctive in using job instead of occupational characteristics.

[5] A 1965 paper by Hodge and Hodge can be seen as the intellectual predecessor to this line of research in sociology. Hodge and Hodge tested what they called the "competition" hypothesis. They argued that if women or blacks gain access to a largely white male occupation, they may be willing to work for lower wages than white men (presumably because their other options are worse), and, thus, employers might lower the wages of all workers in the field. Rather than calling it "devaluation," they called it "competition" because white men were having the wages in "their" occupations driven down by competition from women and blacks. Thus, they argued that, unfortunate for the cause of equality, white male workers have an economic incentive to keep out women and blacks. They relied mainly on cross-sectional regressions. Their only longitudinal analysis showed a negative correlation between 1950 and 1960 change in percent female and change in earnings level in detailed occupations within the major "operatives" category. Occupations with greater increase in percent female lost relative earnings standing. This evidence is consistent with the causal arrow going either way. Although their inference about the role of white male workers in perpetuating discrimination was questioned by Taeuber et al. (1966) (see response by Hodge and Hodge, 1966), the longitudinal correlation was not challenged.

But they also found, supporting the queueing view, that controlling for other factors, the change in mean salary between the periods affected the proportion of women employed in 1983. Their modeling strategy was similar to the lagged-Y-regressor model, except that they controlled for a predicted rather than observed score on the dependent variable in the earlier year, and they made their independent variable a change score. That is, one model took wage level as a function of change in sex composition and earlier predicted wage, and a second model took sex composition as a function of change in wage and earlier predicted sex composition. In our view, it makes more sense to predict level from level or change from change than mix the two in a model. Another weakness of the study is that their unit of analysis was not an administrative job or occupation, but an entire university or college, for which they took an average of all administrative salaries for men and women. Thus, a change in the mix of higher- and lower-paying administrative occupations in a university could not be distinguished from a change in pay level of specific occupations in their analysis.

Baron and Newman (1989) concluded from their study of wage rates in the California Civil Service from 1979 to 1985 that increases in female and minority representation had negative effects on changes in the relative prescribed starting pay of civil service jobs, under stringent controls. The devaluation effect (the negative effect of change in percent female on prescribed wage) was less strong in recently created jobs and in growing lines of work. They did not attempt to estimate an effect of wage on sex composition. Their approach was a lagged-Y regressor model, but they expressed the independent variable as a change score; that is, they estimated the effect of change in sex composition between 1979 and 1985 on 1985 prescribed wage in the job, while controlling for both the 1979 sex composition and the 1979 wage. (As mentioned above, we think it makes more sense to predict a level from a level or a change from a change.)

Snyder and Hudis (1976) used a cross-lag panel model (with lagged-Y regressor). Using US Census Data from 1950, 1960, and 1970, with detailed occupations as cases, they assessed the relationship of sex composition on white males' wages and vice versa (they did not consider women's wages). They found that the proportion female had a negative effect on later male median income, while income did not have a significant effect on an occupation's later proportion female. Thus, their analysis supported the devaluation more than the queueing view (although the article preceded use of these terms in the literature).

Catanzarite (2003) used Current Population Survey data and a panel model to test for pay deterioration in white males' wages in detailed occupations from 1971 to 1981 and from 1982 to 1992. She used a lagged-Y regressor model to assess effects of earlier sex composition on later wages. She found that the earlier proportion white female and the proportion black male in an occupation had a negative effect on later male median income in both time periods. Overall percent female (not broken out by race) had negative effects as well in both periods. She did not test the reverse causal order in this paper, but noted that she did not find the reverse effect (pay affecting sex composition) in unpublished work in progress. Thus, although she is one of the originators of the "relative attractiveness" view, her recent results seem to favor what we have called the devaluation view. She argues, however, that it is not just a matter of cultural values, but that her results can also be interpreted in terms of favored groups (in this case men) having more power to keep their wages from being lowered in a period when wages were falling in many working-class occupations.

Karlin et al. (2002) used the Current Population Survey for 1984–1991, and cross-classified detailed occupation and broad industry to form the cells that were units of analysis. Using various year pairs between 1984 and 1991, they employed cross-lagged panel models (with lagged-Y regressors) to examine the effects of sex composition on average male and female wage and vice versa. They found substantial support for earlier sex composition affecting later wage, but no support for the reverse.

Using Israeli data for 1972–1983, and a lagged-Y-regressor model, Semyonov and Lewin-Epstein (1989) found that occupations higher in percent female had smaller increases or larger decreases in wages for nonminority men, but there was no effect of earlier wage on change in sex composition. Thus, evidence fit devaluation better than queueing theory.

Longitudinal data has also been used to support the institutionalist perspective focused on inertia and path dependence, especially the stability of relative wages. Kim (1999) analyzes the California Civil Service. Wages were first rationalized there in 1931. She quotes a 1930 memorandum in which the Civil Service Commission recommended the following:

> When men and women do the same kind of work to make no difference, but to pay somewhat higher for those occupations filled predominately by men than for those occupations filled predominately by women, where, aside from sex, the qualifications are substantially the same. (Kim, 1999, p. 53, citing a document from the Civil Service Commission from 1930)

Her statistical analysis shows that an occupation's pay in 1931 continued to exert an effect on the pay level it was accorded in the California Civil Service in 1993, even after controlling for 1993 external market wages in the occupations, despite the fact that the state claimed to be using regular market surveys to update wages. Thus, institutional inertia was sufficient to keep in place substantial bias against women's jobs. Kim does not look at whether changing sex composition affected changes in wages or vice versa, as we will here. However, in a model with inertia or path dependence as the central feature, we would not expect these variables to flexibly adjust to each other over time. The point of the model is that the causal processes operating at "birth" or shortly thereafter are permanently frozen in. In this case, "female" jobs were assigned discriminatorily low wages, and inertia kept the relative wages of these occupations low. If Kim's model fits national data, we would expect a continuing cross-sectional relationship between sex composition and wages, but not any adjustment over time of one to the other. And we would expect substantial stability in occupations' relative wages.

Overall, past research shows a fairly robust relationship between occupations' sex composition and wages. Those five studies examining the relationship longitudinally with occupations as units have supported the devaluation view, as has one Israeli study. Only one longitudinal study suggests that wage affects sex composition, and it is limited to academic administration. All these studies have used a cross-lagged panel model with two years (Snyder and Hudis, 1976; Catanzarite, 2003; Karlin et al., 2002) or variations of this model that include the lagged-Y as a control but express the independent variable as a change score (Pfeffer and Davis-Blake, 1987; Baron and Newman, 1989). None have used occupational fixed effects, the model we will argue for here. By using this model we examine whether changes in sex composition around an occupation's long-term average are followed by deviations from an occupation's long term average wage and vice versa. Our contribution here will be to use a longer period (19 years) than prior analyses, to use all the years of data (not just the first and last), and to use what we believe is a superior statistical model for removing omitted variable bias. This should provide the best assessment to date of how occupations' wage and sex composition affect each other over time.

3. Data and methods

3.1. Data

Data for our analyses come from the March ''Current Population Survey: Annual Demographic File'' (called CPS hereafter) for each year from 1983 to 2001 (US Department of Commerce, 1991). Using extracts from the CPS that merged relevant household and family information onto individual records, we selected records to include all civilian workers, aged 16 and older, in the rotation groups that were asked earning questions. Our analysis required computing means, medians, or proportions for each occupation on variables such as earnings and sex composition. In 1983, the CPS began using the 1980 census occupational categories, which were a significant change from the 1970 categories. In 1992, they began using the 1990 categories. The 1990 census made only minor changes to the 1980 categories. Therefore, by combining categories and dropping a few, we were able to construct a set of 3-digit detailed categories that are consistent for the entire period from 1983 to 2001. These approximately 500 Census Occupational categories were our starting point, but we had to collapse these categories to avoid empty cells. Even with a sample size of approximately 40,000 employed individuals per year, no one worked in many of the occupations each year. To solve this problem, we collapsed the occupational categories, combining the small categories in similar fields into larger categories. We came up with 165 collapsed categories that allowed us to have a balanced panel (the same number of occupations for each year), without discarding data from any detailed occupation.[6]

For each of our 165 collapsed occupational categories, for each year, we calculated the sex composition and a sex-specific mean or median on each other variable. We output these cell means, medians, and proportions into files with occupations as the units of analysis. One such data set was created for each year.

3.2. Variables

3.2.1. Log of median wage

March CPS respondents provided their primary occupation and annual salary and wages from the prior calendar year, along with the number of weeks worked in the prior year and the usual hours worked per week in the prior year. We used this information to construct a wage variable for the survey years 1983 to 2001, which measure earnings for the years 1982 to 2000. We constructed wage as the annual earnings (from salary and wages) divided by the product of the number of weeks worked in the prior year and the number of

[6] A list of which detailed 1980 Census occupation categories were put into which of the collapsed category is available upon request from the first author. In a sensitivity test not shown, we repeated all the analyses discussed below with a different strategy—simply discarding all occupations that did not contain at least 50 incumbents in each year. This also yielded 165 occupations, but involved discarding about 20% of employed individuals. Our conclusions from that analysis were virtually identical to those reported here.

usual hours worked per week. The medians were only computed on full-time workers (those who usually worked 35 hours or more per week). In our models, we use the natural logarithm of the median wage of men or women in the occupation in the given year.

3.2.2. Logit of proportion female

For this variable, we start from the proportion female, the number of full-time female employees in an occupation in the given year divided by the number of all full-time employees in that occupation in that year. The variable ranges from 0 to 1. We use this for simple OLS models. Because effects of or on a proportion can be different near the natural limits that the variable can take on (0 and 1), we converted it to the logit of proportion female for our fixed effects models intended to sort out causal order. If we take proportion female of each occupation for each year as p, this is:

$$\text{Logit of proportion female} = \log\left(\frac{p}{1-p}\right)$$

This logged variable has the merit that, given our logarithmic transformation of wage, allows us to have a metric-free way to compare the magnitude of any effect we find of wage on sex composition and any effect of sex composition on wage. We also considered various higher powers of the proportion female or its logit. In addition, to consider the possible nonlinear effects of proportion female another way, we constructed a set of dummy variables for sex composition (*male occupations*, less than 33% female, *mixed occupations*, the reference category, between 33 and 66.9% female, and *female occupations*, at least 67% female).

3.2.3. Control variables

Control variables include sex-specific averages that measure an occupation's male or female workers' characteristics. To measure human capital from learning in school we use average years of education in each occupation per year. To estimate occupational averages of labor force experience (which the CPS does not measure directly), we used potential experience. From individuals' age we subtracted their years of education minus six. Sex-specific averages for each occupation in each year were computed. Many prior cross-sectional analyses testing devaluation have entered control variables measuring skill demands from the Dictionary of Occupational Titles. It is not necessary to include them here because all time-invariant variables are automatically controlled by the fixed-effects analysis.

3.3. Models and estimation

As previously noted, the two most popular ways of adjusting for omitted variable bias in panel models are fixed-effects methods and cross-lagged panel (or "lagged-Y-regressor") methods, that include lagged values of the dependent variable. While earlier studies of devaluation and queueing used lagged-Y-regressor models with only two years, we believe it better to use multiple years of data instead of only two years, since particular years may be idiosyncratic. More important, we believe that, relative to lagged-Y regressor models, fixed-effects modeling does a superior job of removing omitted variable bias (Halaby, 2004). Fixed-effects models deal with omitted variable bias by using only variation within occupation to estimate the parameters. This controls for all stable characteristics of occupations, including those that are not measured.[7] Moreover, in some of our fixed-effects models we also include lagged values of the dependent variable as predictors (the hallmark of the cross-lagged panel approach). We employ a novel estimation method that corrects for bias arising from the possibly reciprocal relationship between wages and proportion female and from the endogeneity of the lagged dependent variable.

Let W_{it} be the natural logarithm of occupations' median wage (with distinct measures for men and women) for occupation i in year t, and let P_{it} be logit of the proportion female for occupation i in year t. We want to estimate models that allow each of these variables to be affected by the other variable in the same year or in some previous year. We also want to control for additional variables represented by the vector X_{it}, which may include lagged variables.

We began with unlagged, *cross-sectional* models estimated separately for each year by ordinary least squares:

$$W_{it} = \beta_0 + \beta_1 P_{it} + \beta_2 X_{it} + \varepsilon_{it}. \tag{1}$$

The sole purpose in estimating these models was to get some sense of the magnitude and direction of the relationship between the two key variables, without regard to reciprocal effects, lags, and other technical issues. For men and women separately, we ran regression models for each year including (sex-specific) education and (sex-specific) potential experience as controls.

Next, we estimated *fixed-effects* models that incorporate reciprocal, lagged effects of the key variables:

$$\begin{aligned} W_{it} &= \beta_0 + \beta_1 P_{i,t-k} + \beta_2 X_{i,t-k} + \alpha_i + \varepsilon_{it}, \\ P_{it} &= \gamma_0 + \gamma_1 W_{i,t-k} + \gamma_2 X_{i,t-k} + \delta_i + \nu_{it}. \end{aligned} \tag{2}$$

In these equations, k is the number of years that the variables are lagged. The disturbance terms ε_{it} and ν_{it} are purely random errors that are assumed to be independent of each other and the vector of X variables. The variables α_i and δ_i represent the effects of all unmeasured variables that vary across occupations but do not vary across time. In fixed-effects models, these variables are allowed to be correlated with all measured time-varying variables.

Our goal was to estimate these two equations using all available years of data, assuming constancy of the regression coefficients across years. But that was not a straightforward task. Because of the reciprocal effects, it is not correct to estimate each equation separately using conventional OLS methods for fixed-effects models (e.g., using dummy variables for occupations or expressing all variables as deviations from occupation means). The reason is that ε_{it} and ν_{it} are necessarily correlated with both P_{it} and W_{it} in later years, violating a key assumption of strict exogeneity (Wooldridge, 2002).

[7] Given that occupational fixed effects only control for unchanging characteristics of occupations, a limitation of our analysis is that occupations' relative skill demands are not entirely stable over time. We lack over-time measures of occupations' skill demands. Nonetheless, we believe that occupational fixed effects are superior to the lagged-Y regressor model in dealing with this problem. The models predicting wage with occupational fixed effects and the lagged dependent variable (Table 1) may provide more control for changing skill levels than the models with occupational fixed effects alone; if changes in skill demands affect wage, then controlling for a recent wage level in effect controls for this.

This assumption can be relaxed by using the machinery of structural equation modeling available in such programs as LISREL, EQS and Amos. (We used the CALIS procedure in SAS.) To do this, the working data set was organized with one record for each occupation (not separate records for each occupation year), with variables having different names for each year of measurement. Although it is possible to estimate the two equations simultaneously, more flexibility can be obtained by estimating them separately.

Consider the first equation with W as the dependent variable, and suppose the lag is $k = 5$ years. (We actually tried and show results from several different lags.) A separate equation is specified for each year of observation from 1987 to 2000 (omitting the first five years because of the five-year lag on the independent variables). The variables have different names in each equation, but the regression coefficients are constrained to have the same values across years. A latent variable, corresponding to the fixed-effect α, is included as an independent variable in the regression. This latent variable is allowed to be correlated with all the measured independent variables in all years. Finally, the error term in each equation (corresponding to ε_{it}) is allowed to be correlated with all *future* values of P_{it} (that is, with proportion female at time t + 1 and beyond). The model is then estimated by maximum likelihood under the assumption that the data are drawn from a multivariate normal distribution. An analogous setup is used to estimate the second equation with P_{it} as the dependent variable. Further details on this approach to estimation can be found in Allison (2005, 134–136).

Fixed-effects models should do an effective job of handling omitted variable bias attributable to time-invariant variables. Nevertheless, to increase our confidence that this problem has been adequately addressed as well as to incorporate approaches commonly used in previous literature, we also estimate models that incorporated both fixed effects and lagged values of the dependent variable. These models have the general form:

$$\begin{aligned} W_{it} &= \beta_0 + \beta_1 P_{i,t-k} + \beta_2 W_{i,t-k} + \beta_3 X_{i,t-k} + \alpha_i + \varepsilon_{it}, \\ P_{it} &= \gamma_0 + \gamma_1 W_{i,t-k} + \gamma_2 P_{i,t-k} + \gamma_3 X_{i,t-k} + \delta_i + v_{it}. \end{aligned} \tag{3}$$

Although models with lagged dependent variables as predictors are known to pose problems for conventional estimation methods (Baltagi, 1995), the structural equation approach described above solves these problems.

4. Results

We start with simple cross-sectional models for each year used in our longitudinal models, 1982 to 2000. We take wage as the dependent variable, although, of course, our subsequent longitudinal analyses are designed to reveal the causal order. Fig. 1 shows the coefficients from proportion female from these cross sectional models for each year. There is a significant negative relationship between an occupation's proportion female and the (natural log of the) median female and median male wages in all years. In these models with controls for education and potential experience, it is significant in all years for both men and women. Since median wage is logged, the magnitude of the coefficients is such that if an occupation's percent female is 10 percentage points higher, its median wage for either men or women is approximately 3–5% less. In the second half of the period, there is some suggestion that the relationship has gotten slightly stronger for men and slightly weaker for women.

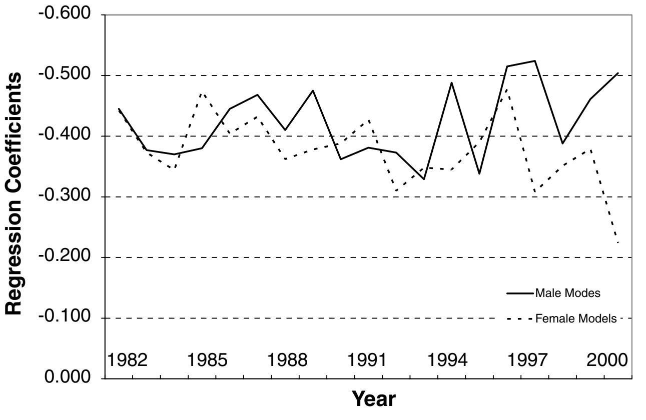

Fig. 1. Coefficients for occupations proportion female from cross-sectional OLS models predicting natural log of male or female median wage, 1982–2000. Note: All coefficients are significant at $p < .05$. All models control for mean (male or female) education and mean (male or female) potential experience.

Alerted to a possible nonlinear shape of this relationship by Cotter et al. (2004), we also tried specifications with proportion female to its second, third, and fourth power. In many years all were significant (results not shown). When we graphed the shape, there was one section around 30–50% where earnings were flat as percent female went up, and a section above 90% female where earnings went up as percent female increased. However, in most of the range of percent female, median earnings went down as percent female went up. So the negative linear shape is a reasonable approximation. In other results not shown, we get similar OLS results when we remove the controls and observe the bivariate relationship, when we add controls for measures of general educational development[8] (GED) and standard vocational preparation (SVP) from the Dictionary of Occupational Titles, or when we take the logit transformation of proportion female.

Table 1 presents our models assessing whether occupational sex composition affects later wage. The table contains coefficients for the logit of proportion female in models predicting later wage. We use fixed-effects models that vary the lag of the independent variable from two to nine years. All models control for education and potential experience. We also present models with and without an explicit control for lagged median (male or female) wage. Inclusion of this control makes the models more similar to the traditional lagged-Y-regressor models, controlling for a fairly recent level of the dependent variable, but they also have the benefit of the fixed effect which can be thought of as controlling for the average wage across all years for the occupation. The devaluation thesis predicts negative coefficients.

What can we conclude from Table 1 about the devaluation thesis? With eight different lags, male and female models, and models with and without controlling lagged wage, there are 32 tests of the hypothesis. Of the 32 tests, just more than half (17) of the coefficients are

[8] Tam (1997) has argued against inclusion of GED in such models because of collinearity with SVP. If we delete this control, we still get significant coefficients for sex composition in most years. Also, we believe that GED should be included (see England et al., 2000 for discussion).

Table 1
Coefficients for lagged logit of occupations' proportion female from fixed-effects models predicting log of later (male or female) median wage, with and without controls for lagged median wage, using pooled longitudinal data, 1982–2000

Effect of logit of proportion female on later log median wage	Female models		Male models	
	FE models w/o lagged D.V.	FE models w/lagged D.V.	FE models w/o lagged D.V.	FE models w/lagged D.V.
2-year lag	−0.0011 (−0.126)	−0.0126 (−1.470)	0.0069 (0.910)	0.0003 (0.042)
3-year lag	0.0134 (1.513)	0.0055 (0.624)	−0.0111 (−1.421)	**−0.0174** (−2.264)*
4-year lag	−0.0082 (0.896)	**−0.0193*** (−2.153)	**−0.0165*** (−2.078)	**−0.0241*** (−3.136)
5-year lag	−0.0183 (−1.890)	**−0.0290*** (−3.061)	−0.0110 (−1.357)	**−0.0177*** (−2.207)
6-year lag	−0.0125 (−1.258)	−0.0150 (−1.541)	**−0.0198*** (−2.359)	**−0.0262*** (−3.256)
7-year lag	−0.0177 (−1.679)	**−0.0308*** (−3.029)	**−0.0174*** (−1.992)	**−0.0315*** (−3.745)
8-year lag	0.0160 (1.436)	−0.0068 (−0.632)	**−0.0193*** (−2.176)	**−0.0408*** (−5.118)
9-year lag	**−0.0210*** (−2.167)	**−0.0206*** (−2.133)	**−0.0225*** (−2.719)	**−0.0237*** (−2.961)

* $p < .05$ (two-tailed test; t statistic in parentheses). All models control for mean (male or female) education and mean (male or female) potential experience. Predicted sign from devaluation hypothesis is negative. Significant coefficients with predicted sign are bolded.

significant and have the predicted sign. (None are significant with the wrong sign.) Thus, there is some evidence for devaluation. Significant effects appear more often in models predicting male than female median wages in an occupation and male coefficients are generally a bit larger. This may be because potential experience taps real experience less well for women than men, and this creates more noise in the female models. It is hard to know a priori exactly how long a shift in sex composition should take to affect wages, but our guess would be that 2 years is too short and 9 years too long; our results show no clear pattern to which lags produce more significant effects. Significant effects occur more often in models controlling for the lagged dependent variable (log of median wage). In one sense such models are more conservative—taking out the occupation's fixed effect on wage (by adjusting around the average wage as all the models do), *and* controlling for the occupation's wage 2–9 years previously (depending on the lag in the model), yet they more often show an effect. Perhaps this is because the lagged Y picks up recent changed conditions in the occupation, and removing this actually takes out noise that otherwise obscures the tendency for occupations to lose relative wage as they feminize. Although there are quite a few models, at least for men, in which coefficients are significant, effect sizes are very small in magnitude. Most coefficients are in the neighborhood of −.02; this means that every increase of one percentage point in percent female leads to an approximately .02 of one percent decrease in median wage in the occupation.[9]

As another way to test devaluation, consider the specification in Appendix Table 1 (female models) and Appendix Table 2 (male models). These models differ from those in Table 1 only in that they enter dummy variables representing a trichotomous version of percent female of the occupation. As such, they allow nonlinearity. The prediction is that the effect of an occupation being female relative to mixed is negative, while being male relative to a mixed occupation is positive. Appendix Table 1 shows virtually no support for

[9] If we enter higher order terms for the logit of proportion female in the models in Table 1, unlike the OLS models using proportion female in Fig. 1, these higher order terms are seldom significant in the fixed effects models (results not shown).

devaluation when women's wages are the dependent variable (3 out of 32 tests show the predicted sign with significance). In Appendix Table 2 where we predict male median wages, and examine effects of an occupation being female as opposed to mixed, tests for half (4 out of 8) of the lag lengths are significant when lagged wage is explicitly controlled. (But effect of male occupation relative to mixed is never significant, and models without the lagged dependent variable show only one significant effect.) The magnitude of these effects is not entirely trivial. Coefficients are in the −.03 to −.05 range, indicating that if an occupation goes from being mixed (33–66% female) to female (67% or more female), median male wages go down by 3–5%. (The difference between male and female occupations is larger than this, although the contrast between mixed and male is never significant in the male equations.)

Thus, overall, considering Table 1 and Appendix Tables 1 and 2, there is some support that, when occupations feminize, at least male wages go down, although evidence is mixed, effect sizes are quite small, and there are few significant effects on women's wages.

Table 2 presents our fixed-effects models aimed at testing the queueing or relative attractiveness theories, which state that earlier wage should affect later sex composition. The predicted effect is negative—that is, increases in wage should lower percent female, and decreases in wage should increase percent female. Again, we vary the lag, and whether we explicitly control for the lagged dependent variable (in this case, the logit of proportion female). There is absolutely no support for this hypothesis. Out of 32 coefficients, none have the predicted significant negative sign. Nor did removing the education and potential experience controls change things; in no case did we find any of the predicted negative effects (results not shown).

In other analyses not shown we tested separate parts of the queueing thesis with fixed effects models: that earlier wage affects later number of men (the predicted sign is positive) and also affects later number of women (with a predicted negative sign). These predictions were not upheld either; none of the relevant coefficients were significant. In short, there is no evidence in this 19 year period that reductions in occupations' relative wages led to their feminization.

Table 2
Coefficients for lagged log median (female or male) wage from fixed-effects models predicting logit of proportion female, with and without controls for lagged logit of proportion female, using pooled longitudinal data, 1982–2000

Effect of on log median wage on later logit of proportion female	Female models		Male models	
	FE models w/o lagged D.V.	FE models w/ lagged D.V.	FE models w/o lagged D.V.	FE models w/ lagged D.V.
2-year lag	0.0070 (0.161)	−0.0143 (−0.333)	0.0599 (1.191)	0.0594 (1.179)
3-year lag	−0.0152 (−0.354)	−0.0324 (−0.754)	0.0523 (−1.041)	0.0439 (0.880)
4-year lag	0.0419 (0.932)	0.0212 (0.470)	0.0981 (1.912)	0.087 (1.689)
5-year lag	0.0563 (1.189)	0.0363 (0.767)	0.0127 (0.235)	−0.0065 (−0.120)
6-year lag	0.0471 (0.994)	0.0285 (0.600)	0.0391 (0.702)	0.0171 (0.305)
7-year lag	0.0557 (1.151)	0.0366 (0.752)	−0.0041 (−0.071)	−0.0238 (−0.409)
8-year lag	0.0361 (0.721)	0.0162 (0.320)	−0.0593 (−0.981)	−0.0913 (−1.501)
9-year lag	0.0188 (0.355)	0.0170 (0.319)	0.0378 (0.601)	0.0191 (0.303)

Note: All models control for mean (male or female) education and mean (male or female) potential experience. Predicted sign from queueing or relative attractiveness hypothesis is negative. None of the coefficients are significant.

5. Conclusion

There is clearly an association between occupations' sex composition and their wages, as many past cross-sectional studies have found, and as we show in Fig. 1. More female occupations pay less. At issue in this paper is how the over-time causal dynamics between these two factors work, and what statistical model is most appropriate for assessing longitudinal causal dynamics. Past longitudinal research on how wages and sex composition co-vary over time has come from various versions of cross-lagged panel (lagged-Y regressor) models, and generally supported devaluation but not the queueing or relative attractiveness view. Past US studies by Snyder and Hudis (1976); Pfeffer and Davis-Blake (1987); Baron and Newman (1989); Karlin et al. (2002); Catanzarite (2003) all supported the longitudinal implication of the devaluation hypothesis. So did Semyonov and Lewin-Epstein (1989), using Israeli data. The only study finding an effect running from earlier wage to later sex composition is that of Pfeffer and Davis-Blake (1987), who use institutions as units of analysis (i.e. taking average sex composition and average salaries of all administrators in a university, rather than occupations or jobs within institutions). All these studies used pairs of years and a cross-lag panel approach, with the lagged measure of the dependent variable the main antidote to omitted variable bias. In results not shown, we performed cross-lag panel models on these data and got similar results—a fair amount of support for devaluation and none for queueing/relative attractiveness.[10] Given the past studies, and our preliminary results from cross-lagged panel models, when we began our fixed-effects modeling we thought that we too would find support for the longitudinal implications of the devaluation view but not for the queueing or relative attractiveness view. Using our more stringent fixed effects models, we did find some limited support for the idea that feminization lowers wages, although the effects were quite small, and they appeared mainly in models predicting men's (not women's) median wage.

Our fixed effects models unearthed absolutely no evidence in support of queueing or relative attractiveness; coefficients were never significant as predicted in any specification we tried. There was simply no evidence, regardless of the lag, that occupations with declining relative wages (net of the education and potential experience of their incumbents) feminized more than other occupations.

We engaged in this project because of our belief in the superiority of an appropriate fixed-effects approach over the traditional cross-lagged panel approach. (On the merits of fixed-effects, see Halaby, 2004; on the limits of the cross-lagged panel approach, see Allison, 1990.) One way of thinking of the difference between these two approaches is that fixed-effects models use only variation within each occupation, and thus control for all stable characteristics of occupations. We also thought it important to pool multiple years of data, which is possible within the lagged-Y-regressor approach, but had not been done by any past studies on the question.

Our fixed-effects analyses show no causal effect of change in wage on change in sex composition (as predicted by queueing and relative attractiveness hypotheses), and some but

[10] With our data we created two year pairs (1982 and 1991, 1991 and 2000) and ran cross-lagged panel models predicting sex composition from lagged (male or female median) wage, and then predicting wage from lagged sex composition. The lagged dependent variable was always included. We found no evidence that wage affects later sex composition in the direction predicted by the queueing view (or that it affects later number of men or women as predicted). We found some evidence that early sex composition affected later wage in these models.

very modest evidence of changes in sex composition affecting changes in wage (as posited by devaluation theories). Why did past studies without fixed effects find clearer evidence than we did that changes in sex composition lead to changes in wages? We suspect it is because they had inadequate protection against omitted variable bias. We return below to what we think the omitted variable is.

How do we interpret this result in the broader theoretical framework motivating our analysis? One interpretation is consistent with the perspective of those who have argued that there is some advantage of female occupations that leads to their low wages (Filer, 1985; Tam, 1997; Macpherson and Hirsch, 1995). These authors propose theoretical reasons not to expect a causal relationship; the reasons can be subsumed under the theory of equalizing differences. Tam argued that women's occupations require less occupational-specific training, obtained on or off the job, and that the market rewards occupation-specific more than general training. Economists have argued that there must be something attractive about the nonpecuniary features of predominantly female occupations (e.g. mother-friendliness) that leads women to be willing to enter them for lower wages. So one interpretation of our very limited support for devaluation is that these theorists are right—the lower pay of female occupations is (with the exception of the small effects we found) all compensating differentials or unmeasured human capital demands. Those could be the omitted variables. This is not the only possible interpretation, however, and not the one we find most plausible.

As mentioned in our introduction, a second theoretical perspective—the institutionalist view—predicts inertia and lack of flexible responses of wage to sex composition or vice versa. Our conjecture is that the relationship *was* once causal in the beginning stages of the development of occupations. It might have worked something like this: As occupations new to the economy or to a certain locale were created, if, for whatever reason, an occupation was filled largely with women, it was assigned a lower wage because of this fact, as the devaluation perspective argues. Often occupations would attract and employers would seek women because the task was stereotyped as "female" and thus employers thought women more appropriate. In these cases, we speculate that they set wages lower both because women were thought to need and deserve lower wages than men (recall that prior to 1963, paying women less than men even in the *same* job was legal and well accepted) and because the cultural devaluation of women had "rubbed off" onto female-typed tasks. This would be consistent with Kim's (1999) "smoking gun" memo from the early California Civil Service in which officials were instructed to set lower wages in occupations filled by women than occupations filled by men requiring equal credentials. Kim notes that the same consultant used by the California Civil Service designed many other public and private wage systems around this time.

At these origin points we also conjecture that the causal arrow ran the other way. Lower paid occupations probably were more open to women in eras of legal and widely consensual hiring discrimination, and this would have led wages to affect sex composition, as the queueing and relative attractiveness views suggest. The hiring discrimination underlying this conjecture seems a reasonable assumption given that prohibiting hiring discrimination by sex was seen as laughable when first proposed as an amendment to the Civil Rights Act of 1964 (by a Southern Congressman seeking to defeat the bill's main goal of equal opportunity by race).

In this institutionalist view, the persisting correlation between sex composition and wage is mostly because of inertia in both variables, particularly in relative wages. This inertia in each means that movement in the other variable is not flexibly responded to in either causal direction. In this view, gender bias in either hiring or valuing occupations may be causal, but most of the causation has its effect at "birth," rather than in ongoing dynamics. In this view, the cross-sectional relationship reflects past rather than ongoing causal dynamics. Of course, our present analyses cannot help us assess whether these historical conjectures about causal effects at the birth of occupations and organizations are valid. But they are consistent with the possibility that, while there is some change in both relative sex composition and relative wages of occupations, the fundamental fact is the stability of both (at least relative to changes in the overall wage level and overall proportion of women in the labor market). Indeed, in our data the correlation between occupations' 1983 and 2001 percent female is .95.[11] The correlation between an occupation's median male wage in the two years is .82, with a .78 correlation between occupations' median female wage in the two years. The original causal effect of one on the other may now live on mostly because of this inertia. Thus our best guess is that an originally truly causal relationship between sex composition and pay, combined with ongoing institutional inertia, are the main reasons for the enduring association between sex composition and wage found in so many studies.

What our study has shown, if we believe the fixed-effects models, is that the *dynamics* of occupations' sex composition are not strongly related to the *dynamics* of their wage. If this is true, then our longitudinal analysis is not the appropriate way to see how underpaid female occupations are, because most of this underpayment is a response to the original sex composition of the occupation with little flexibility afterwards. Since the rank order of occupations' sex compositions has changed little, much of this underpayment is preserved, but is stripped out of our fixed-effects estimates. If this is true, what it implies for devaluation theory is that, using skill levels as a barometer of "worth," women's occupations are underpaid relative to men's, as England (1992) and Steinberg (2001) and others have argued. But, the longitudinal version of the hypothesis—that feminization would lead to falling (or less rapidly growing) wages—appears to have much weaker support. Put most simply, jobs that start out as female are quite seriously underpaid, but when jobs increase their percent female, this has little effect on their wage.

What are the implications of our analysis for the queueing (Reskin and Roos, 1990) or relative attractiveness (Strober and Catanzarite, 1994) thesis? Authors offering this view have presented compelling case studies suggesting that when occupations downgrade in pay or status they feminize (Reskin and Roos, 1990; Strober, 1984; Strober and Arnold, 1987). But we find no support for this in our fixed-effects results, and all but one (Pfeffer and Davis-Blake, 1987, on college administrators) past cross-lagged panel studies were disconfirming as well. This suggests that their case studies do not generalize, at least in the period after 1983. That is, in this period, women have entered occupations improving their skill-adjusted wages as much as they have entered those with declining relative wages. This does not mean that there is no sex discrimination in hiring. It undoubtedly exists, but the results do seem inconsistent with the view that employers prefer men in all occupations, one of the assumptions used to generate predictions in the queueing view. We believe that a more realistic model of hiring discrimination is that, due to stereotypes, employers prefer to hire men in male occupations and prefer to hire women in female occupations.

[11] On demographic inertia in sex composition in academia and its effects, see Hargens and Long (2002).

Overall, our analysis suggests that an important enemy of gender equality is the inertial relationship between highly female occupations and low wages, a kind of "original sin" in the system which disappears only glacially without active policy intervention, and contributes to the sex gap in pay as long as there is substantial segregation.

Appendix A

See Appendix Tables 1 and 2.

Appendix Table 1
Coefficients for dummy variables representing lagged proportion female from fixed-effects models predicting log of later ***female*** median wage, with and without controls for lagged median wage, using pooled longitudinal data, 1982–2000

Effect of lagged female or male occupation (relative to mixed) on later log female median wage	Models predicting female median wage			
	Without lagged D.V.		With lagged D.V.	
	Female occupation	Male occupation	Female occupation	Male occupation
2-year lag	0.0028 (0.194)	−0.0271 (−1.803)	−0.0069 (−0.482)	−0.0230 (−1.531)
3-year lag	−0.0058 (−0.399)	0.0057 (0.375)	−0.0138 (−0.960)	0.0091 (0.599)
4-year lag	0.0031 (0.212)	**0.0372*** (2.352)	−0.0049 (−0.328)	**0.0414*** (2.645)
5-year lag	−0.0221 (−1.408)	0.0150 (0.890)	−0.0300 (−1.918)	0.0217 (1.299)
6-year lag	−0.0313 (−1.921)	−0.0316 (−1.861)	**−0.0402*** (−2.473)	−0.0239 (−1.412)
7-year lag	−0.0083 (−0.459)	0.0067 (0.358)	−0.0193 (−1.074)	0.0151 (0.812)
8-year lag	0.0101 (0.549)	0.0134 (0.692)	−0.0016 (−0.088)	0.0263 (1.385)
9-year lag	0.0014 (0.072)	−0.0130 (−0.630)	−0.0146 (−0.745)	0.0021 (0.102)

* $p < .05$ (two-tailed test; t statistic in parentheses). All models control for mean (male or female) education and mean (male or female) potential experience. Occupations from 67 to 100% female are "female"; those from 0 to 33% female are "male." Others are in the reference category, "mixed." Predicted sign from devaluation hypothesis is negative for female occupations and positive for male occupation. Significant coefficients with predicted sign are bolded.

Appendix Table 2
Coefficients for dummy variables representing lagged proportion female from fixed-effects models predicting log of later ***male*** median wage, with and without controls for lagged median wage, using pooled longitudinal data, 1982–2000

Effect of lagged female or male occupation (relative to mixed) on later log male median wage	Models predicting male median wage			
	Without lag D.V.		With lag D.V.	
	Female occupation	Male occupation	Female occupation	Male occupation
2-year lag	−0.0253 (−1.768)	−0.0112 (−0.713)	**−0.0302*** (−2.120)	−0.0088 (−0.557)
3-year lag	−0.0237 (−1.635)	−0.0035 (−0.219)	**−0.0292*** (−2.010)	0.0011 (0.069)
4-year lag	−0.0114 (−0.767)	−0.0228 (−1.395)	−0.0172 (−1.170)	−0.0184 (−1.136)
5-year lag	**−0.0469*** (−2.961)	−0.0004 (−0.022)	**−0.0504*** (−3.204)	0.0046 (0.268)
6-year lag	−0.0094 (−0.578)	0.0173 (0.992)	−0.0201 (−1.235)	0.0267 (1.539)
7-year lag	−0.0329 (−1.944)	−0.0009 (−0.053)	**−0.0416*** (−2.462)	0.0109 (0.599)
8-year lag	−0.0104 (−0.600)	−0.0174 (−0.941)	−0.0297 (−1.729)	0.0015 (0.085)
9-year lag	0.0029 (0.155)	0.0210 (1.081)	−0.0116 −0.629	0.0360 (1.861)

* $p < .05$ (two-tailed test; t statistic in parentheses). All models control for mean (male or female) education and mean (male or female) potential experience. Occupations from 67 to 100% female are "female"; those from 0 to 33% female are "male." Others are in the reference category, "mixed." Predicted sign from devaluation hypothesis is negative for female occupations and positive for male occupation. Significant coefficients with predicted sign are bolded.

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