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OCCUPATIONAL GENDER COMPOSITION AND WAGES IN CANADA: 1987-1988

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ABSTRACT

The relationship between occupational gender composition and wages is the basis of pay equity/comparable worth legislation. A number of previous studies have examined this relationship in US data, identifying some of the determinants of low wages in ``female jobs'', as well as important limitations of public policy in this area. There is little evidence, however, from other jurisdictions. This omission is particularly disturbing in the case of Canada, which now has some of the most extensive pay equity legislation in the world. In this paper, we provide a comprehensive picture, circa the late 1980's, of the occupational gender segregation in Canada and its consequences for wages. The sample period precedes many provincial pay equity initiatives and thus the results should provide a baseline for the evaluation of this legislation. We find that the estimated wage penalties in female jobs in Canada are generally much smaller than the estimates for the United States. Although there is some heterogeneity across worker groups, on average, the link between female wages and gender composition is small and not statistically significant.

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1. Introduction

Canada has some of the most extensive comparable worth/pay equity legislation in the world. It covers public sector workers at the federal level and in most provinces. Recently pro-active policies have been extended to the private sectors in Ontario and Quebec, provinces which contained 62 percent of the population of Canada in 1996.[2] One might infer that the source of this legislative activity is a battery of studies documenting systematically lower wages in "female jobs". In fact, this is not the case. While there is a large literature documenting the gender wage gap in this country, there is, to our knowledge, next to no evidence that female jobs are poorly paid in Canada.[3] There is extensive evidence, however, of a negative effect of the "femaleness" of occupations on wages in other parts of the world, particularly the United States. The basis of the legislative initiatives, therefore, would appear to be the experiences of other countries.

The evidence from the United States provides important information on the determinants of low wages in "female jobs", as well as limitations of public policy in this area. For example, Johnson and Solon (1986) show that male-female wage differences within occupations are primarily inter-firm/industry and thus outside the purview of most comparable worth legislation.[4] Macpherson and Hirsch (1995) argue that much of the correlation of wages with the femaleness of occupations can be accounted for by detailed job characteristics, and differences in the unobserved skills of workers in male and female jobs. That said, advances in other research areas suggest that these sorts of inferences can be enriched

[2]Source: 1996 Canadian census. See http://www.statcan.ca/english/census96/.

[3]See, for example, Baker, Benjamin, Desaulniers and Grant (1993), for references on the analysis of the gender gap in Canada. Baker et al. (1993) attempt to estimate the correlation of wages with the femaleness of employment in Canada as of 1985. Their analysis is limited by the lack of appropriate occupational data. Fillmore (1990) finds a very small effect of percentage female on average occupational female earnings. Finally, Reilly and Wirjanto (1999) find a substantial penalty to gender composition of the establishment.

[4]This is consistent with studies such as Carrington and Troske (1995) that document gender segregation across firms and the role it plays in the gender wage gap.

by corroborating or contrasting evidence from other jurisdictions. In the case at hand, however, the empirical evidence is overwhelmingly drawn from the U.S. experience. An investigation of the Canadian data, therefore, widens the empirical perspective, and brings a new set of facts to bear on theories of wages in female jobs. Canadian evidence may be particularly enlightening in comparison to U.S. results, because of similarities in the culture, labour market institutions and economic trends in the two countries.

In this paper we provide a comprehensive picture, circa the late 1980s, of occupational gender segregation in Canada and its consequences for wages. We examine both conventional estimates of the relationship between the femaleness of occupations and wage rates, and alternative representations of the relative position of female jobs, such as kernel density estimates. We also document the heterogeneity in the status of female jobs across workers distinguished by full/part time and union status, as well as other demographic characteristics.

Our choice of time period is deliberate. In the late 1980s, there were few public sector, and no private sector, pay equity initiatives in Canada. The labour market was largely free of the effects of comparable worth. In contrast, by the late 1990s pay equity is almost synonymous with the public sector, and is soon to cover the private sectors of Canada's two largest provincial labour markets. Evaluation of this flurry of legislative activity requires careful documentation of the wages in female jobs *before* these laws took effect. Estimates of the penalty to work in female jobs from this period provide, by some measures, an upper bound on the potential benefits of these initiatives. Therefore, the analysis we offer in this paper can serve as a baseline for future work in the area.

Our investigation uncovers some surprising differences in the Canadian case. Relative to recent evidence of the wage penalties for female jobs in the United States, the penalties in Canada, for both males and females are smaller. Furthermore, our estimates of the penalty for Canadian females are consistently small and not significantly different from zero.

We begin in Section 2 surveying the legislative environment in Canada at the time of the analysis. The description of the data and its salient features are presented in Section 3. Section 4 outlines our econometric strategy for estimating the correlation of occupational gender composition and wages in the presence of grouped data. The results are presented in Section 5. They reveal that the link between female wages and gender composition in Canada is generally small and not statistically significant. In Section 6 we examine the relationship between the "wage penalties" in female jobs and the gender wage gap. We conclude in Section 7.

2. The Legislative Environment

The objective of pay equity legislation is to eliminate the effect of occupational segregation by gender on wages. Empirically, this means the elimination of any systematic relationship between wages and the femaleness of employment, net of differences in "allowable" productivity related characteristics across individuals in different occupations.[5] This relationship is the primary focus of the study. While a comprehensive summary of pay equity in Canada is beyond the scope of this paper, it is necessary to consider the pay equity policies in effect in Canada at the time of our analysis (1987 and 1988). These policies have obvious implications for the interpretation of wage levels we observe in female jobs.

Canada has been called a world leader in comparable worth (e.g., Weiner and Gunderson (1990)).[6] That said, in our period of interest many provincial pay equity initiatives were quite recent, and should have had limited effects in the labour market. Two of the longer

[5]Some studies, such as Blau and Beller (1988), investigate the relationship between the femaleness of employment and wages using dummy variables for male dominated employment and mixed employment. Yet other studies (Killingsworth 1990) combine dummy variables with percentage female. We focus on "percentage female" for comparability with the more recent studies.

[6]Good summaries of the state of Canadian legislation around our sample period can be found in Symes (1990) and Weiner and Gunderson (1990). The current legislative environment is summarized in CCH Canadian Limited (1997).

standing policies in place in the late 1980s were in Quebec and in the federal sector. The concept of pay equity was introduced to the human rights codes of these jurisdictions in 1977 and 1978, respectively. Both of these pay equity initiatives were complaint-based. Under complaint based legislation, investigation of (and possible restitution for) low wages in female jobs is only initiated if an employee complaint is registered. Therefore, the onus is on workers. The alternative is a proactive program, in which employers erect a pay equity plan that typically involves 1) the identification of predominantly female and predominantly male jobs, 2) the assignment of numerical scores to jobs reflecting their levels of skill, effort, responsibility, and the working conditions, 3) the comparison of the numerical scores of female and male jobs in relation to salary rates, and 4) pay adjustments for 'undervalued' female jobs. Here the onus is on the employers.

The Quebec legislation in principle covered all employees in the province working outside the federal jurisdiction. This seemingly wide ranging legislation was rarely used, however, with only 37 cases heard by 1990 (Weiner and Gunderson 1990). The federal legislation covers both the (broader) federal public sector and federally regulated industries (e.g. transportation, banking).[7] It also appears, however, to have been seldom used in the period preceding our years of interest. By 1990 roughly 20 cases, affecting just 5000 workers, had been heard under the legislation (Weiner and Gunderson 1990).[8]

Pay equity in other jurisdictions circa the late 1980's was quite recent and typically restricted to the public sector. Manitoba passed the first pro-active pay equity legislation in 1985. The first awards were to be made by September 1987, which is one of our sample years. Since the implementation of this legislation proceeded on schedule, it is possible that its initial effects, if any, will be captured in our data. The next initiatives were in Ontario

[7]These also include crown corporations.

[8]See Symes (1990) and Cihon (1988) for further evidence that the federal and Quebec pay equity legislation of this period was seldom tested.

in 1987 and in Nova Scotia and Prince Edward Island in 1988.[9] The implementation plans for this legislation suggest that their effects are likely outside our sample period.[10]

Therefore, in the late 1980's Canada's labour market might be considered largely free of any effects of pay equity policies, save for the rarely used federal and Quebec laws, and any initial effects of Manitoba's legislation.[11] Of course, the 1990s have seen a flurry of initiatives such as the implementation of pro-active pay equity in the private sector in Ontario, and more recently in Quebec.[12] By choosing our sample period prior to this legislation, we are able to isolate its target.

3. Data and Descriptive Evidence

The data for this study are drawn from the Labour Market Activity Survey (LMAS) for 1987 and 1988. We include all wage and salary workers between the ages of 16 and 69, who are not full-time students and are earning more than $1.00 an hour.[13] As explained below, additional variables measuring gender composition are obtained from Census data

[9]Newfoundland had a non-legislated pay equity initiative as of 1988.

[10]For example, the first awards under the Ontario legislation were scheduled for January 1, 1990.

[11]It is possible that the threat effect of the Quebec and federal legislation led some firms in these jurisdictions to change their pay structures. While we lack the data to examine the evolution of the effect of gender composition in different jurisdictions over the 1980's, we can examine any provincial heterogeneity in that effect as of 1987/88. Our analysis by provinces for 1987 and 1988 combined (to get larger sample sizes) reveals that the effect of the occupational femaleness rate, $PFEM$, on female wages is generally small and not statistically significant ranging from -0.051 to 0.113 with standard errors around 0.06. The signs of the coefficients are not obviously related to the existence or forthcoming implementation of provincial pay equity legislation: Newfoundland (-0.021), Nova Scotia (0.113), New Brunswick (-0.009), Quebec (-0.051), Ontario (-0.040), Manitoba (-0.001), Saskatchewan (0.094) Alberta (0.018), British Columbia (0.048).

[12]The first pay equity awards in the Ontario private sector were scheduled for January 1, 1991, while the Quebec legislation passed in 1996 will not be implemented until 2001.

[13]We exclude full-time students because they are excluded under most legislation, when they work in connection to their studies. This exclusion is also made for comparability with other studies (Macpherson and Hirsch 1995).

and variables measuring occupational characteristics are coded from the Canadian Classification and Dictionary of Occupations (CCDO). The LMAS is a retrospective survey covering year-round labour market activity. To mimic a point-in-time survey, we select job information as of the third week of November.[14] Wages are obtained from the main job at this time; they are the actual hourly wage for workers paid by the hour and the usual hourly earnings for other workers. The resulting sample sizes are given in table 1, which also provides the average wage levels in 1988 Canadian dollars by gender.[15]

We measure the femaleness of occupations as the proportion of employment which is female in the corresponding four-digit detailed occupation ($PFEM$). To minimize measurement error, $PFEM$ is constructed from the 1991 Canadian census (the reference year is 1990).[16] In each case, we sample individuals who are employed in the reference week and otherwise satisfy the same selection criteria as for the LMAS.[17] The Canadian four-digit occupation codes, comprising approximately 500 categories, are roughly of the same order of aggregation as U.S. three-digit codes.[18]

We note that an evaluation of the Canadian evidence has not been possible in the past because public use data sets include coarse occupation codes. For example, Baker et al. (1993) provide some evidence of the correlation of wages with the femaleness of employment in Canada as of 1985. Their results, however, are from *Survey of Consumer Finance* data in which occupation is available at only the 2-digit level (i.e., 47 categories). Furthermore,

[14]That particular choice of week was dictated by comparability with other surveys in the context of a larger research project.

[15]The LMAS provides sample weights that are used in the analysis.

[16]The Canadian 1980 SOC occupational codes available in the LMAS are also available in the 1991 census.

[17]For example, we exclude individuals from the Yukon and Northwest Territories from the Canadian Census since they are not surveyed in the LMAS.

[18]There are, however, some differences in the coding of occupations across the two countries that should be considered in any attempt to compare the results of this study to those in the U.S. literature. In Baker and Fortin (1999c) we provide a Canada/United States comparison of the wage penalty in female job's using a cross-country consistent set of occupation codes.

they demonstrate that estimates of the correlation are sensitive to the aggregation of the occupational categories.[19] We were fortunate to gain access to versions of the census and LMAS files that include the more detailed occupation codes.[20]

In table 1 we provide an overview of the gender composition of occupations and its consequences for wages in Canada in 1987 and 1988. Across all jobs, the femaleness rate, $PFEM$, is about 67 percent for women, while for men it is 25 percent. As a point of comparison, we also report corresponding statistics for the United States. These are calculated from the Outgoing Rotation Group samples of the Current Population Survey (CPS-ORG) for 1987 and 1988, using similar sample exclusions and U.S. 3-digit occupation codes. The average $PFEM$ in the United States is very similar to the rates observed in Canada, as well as to the rates reported in Macpherson and Hirsch (1995) for these years. The statistics are also reported for "female", "mixed" and "male" jobs. Predominantly female jobs are defined as those with a femaleness rate of 60 percent or higher.[21] In 1988, they represented 57 percent of female employment in Canada and 61 percent in the United States. Clerical and health care work are typical female jobs. Predominantly male jobs are those with a femaleness rate of at most 30 percent. In 1988, they represented 9.8 percent of female employment in Canada and 8.5 in the United States. Truck driving and mechanical repair are typical male jobs. Other jobs are mixed. In 1988, they represented 33 percent of female employment in Canada and 30 percent in the United States. Managerial jobs and work in food preparation and processing are typical mixed jobs. Again $PFEM$ is very similar in

[19]They compare estimates of the correlation of wages with the gender composition of employment in SCF data using, alternatively, 1-digit (i.e., Canadian Census) and 2-digit occupational codes. The correlation's for females are 0.354 (0.028) and 0.055 (0.034) from the 1-digit and 2-digit codes respectively (standard errors in parentheses). Similar changes are reported for the correlation's for males.

[20]In addition to detailed occupation codes, our Canadian data also contain a single year age variable instead of the usual 5–year classes available in the LMAS.

[21]These definitions of male and female jobs are the more recently used in actual legislation's, in the Ontario *Pay Equity Act.*, for example.

the two countries in this decomposition. The Duncan index is a convenient summary of this information, and it confirms the similarity of occupational gender composition in the two countries: it is equal to 59 percent in Canada and 58 percent in the United States.[22]

We also report average wages (in 1988 Canadian dollars) for all jobs and average wages by job type. Here we see some interesting Canada/U.S. differences. In the United States, women in female jobs are the lowest paid on average while women in mixed jobs are the highest paid. In Canada, it is the women in mixed jobs who are the lowest paid. None of the differences in average wages across job types would be statistically significant given the large standard deviations, but these descriptive statistics give a flavour of the results to come.

The corresponding unadjusted female/male wage ratio is also reported in the last column of the table. The ratio averages 76 percent in Canada (for all jobs) and 72 percent in the United States; consistently higher in Canada, although the cross-country difference is not substantial.[23] These ratios are higher then those typically reported for full-time full-year workers (approximately 0.65 for Canada in 1988). We argue that selecting full-time full-year workers introduces a different selection bias among men than among women. Excluding part-timers and seasonal workers among men throws out workers who are more marginally attached to the labour market leaving a wage distribution more skewed to the left. Because many women choose to work part-time or part-year for family reasons, these part-timers are more evenly distributed across the entire female distribution. Their exclusion does not distort the wage distribution as much as it does for males. To account for the fact that more women than men work part-time, a more appropriate correction is to weight the data by hours of work. This correction actually raises the female/male wage

[22]The Duncan index of segregation provides a measure of the concentration of women in certain occupations. Recall that this index can be interpreted as the proportion of the male or female employed population that would need to change occupations to achieve an even distribution.

[23]Macpherson and Hirsch (1995) report unadjusted female/male wage ratios for the U.S. of 0.692 for 1987 and 0.699 for 1988.

ratio by about 1 percentage point in both countries.

We also report the estimated coefficient $\widehat{\gamma}$ from the regression, $\ln w_i = \delta + \gamma\, PFEM_i + \epsilon_i$, estimated by weighted least-squares, using LMAS and CPS-ORG sample weights respectively. It is not surprising that for Canada the estimate of γ for females is effectively 0, as, on average, the lowest paid women are in mixed rather than female jobs. In contrast, for U.S. women the implied elasticity at an average percentage female of 0.67 is (0.67×-0.227) -0.152. The two countries switch places in the comparison for male. It is now the American men that face the much smaller penalty.

The occupations "driving" the simple regression coefficients for Canadian women and men are illustrated in figures 1 and 2 respectively. We plot the regression line of average occupational log wages on the femaleness rate of the occupation. The relative sizes of the circles indicate the relative weights of the occupations. These pictures clearly show that the regression line is essentially flat for women, while it is negatively sloped for men. For women, a sizeable number of relatively highly paid nurses and kindergarten and elementary teachers would seem to compensate for an equal number of relatively low paid waitresses and cashiers. Among men, there is a sizeable proportion of the workforce employed as waiters and cashiers but not as many nurses and teachers.

In figure 3 we plot weighted kernel regressions of the relationship between wages and femaleness for both women and men.[24] Both lines reveal some non–linearities, in particular a dip around the 55 percent femaleness rate; this is attributable to the sales clerk occupation which is typically paid the minimum wage. Moving above the 70 percent rate females' wages appear to rise, or at least not fall, while males' wages appear to decline. This figure also brings to light an interesting aspect of the gender wage differential: it is present at all

[24]Kernel regressions are easily understood with reference to moving averages. Around any femaleness rate, a moving average could be computed as the sum of average occupational wages times a rectangular weighing function of a given width. The corresponding kernel regression would be computed as the sum of average occupational wages times a Gaussian weighing function, called the kernel, of given bandwidth. Here, the bandwidth used is 0.05.

levels of the femaleness rate, and is larger in male jobs.

Differences in the wages of women and men may result from differences in individual characteristics, in particular in human capital variables. In table 2 we report the average characteristics of the Canadian samples by gender. The table shows that on average working women appear to be more educated than working men. Approximately the same proportion of women and men hold a university degree, while a higher proportion of women hold a post-secondary degree and a lower proportion are drop-outs. On the other hand, there are other gender differences that work to the disadvantage of women. They have lower levels of tenure and work in smaller firms.

Another difference between working men and women is the proportion of workers covered by collective bargaining. In our samples, the union coverage rate for men is 8 points higher than for women. Doiron and Riddell (1994) argue that the gender wage gap would have increased by 7 percentage points between 1981 and 1988 if not for the reduction in the gender unionization gap which occurred over this period. [25] An illustration of the potential impact of unionization on the effect of gender composition on women's wages is shown in figure 4. We plot kernel density estimates, which can be understood as smoothed histograms, of wages by job type for both women and men.[26] The union coverage rates

[25]Our rates are, as well as those of Lemieux (1993) and Riddell (1993), are higher than those reported by Doiron and Riddell (1994) for 1988 LMAS (38 percent for males and 29 percent for females). Based on the LMAS, Riddell (1993) reports (p. 113) union coverage rates of 43.7 (40.5) percent for males and 35.2 (34.3) percent for females in 1986 (1990). Lemieux (1993) who uses the merged 1986-87 LMAS longitudinal files, reports (p. 76) union coverage rates of 45.8 percent for males and 36.4 percent for females. In addition to any effects of the differences in survey years, part of the difference appears to be due to our exclusion of full-time students. Adding these individuals back into our sample we obtain unionization rates of 43.2 percent for males and 35.4 percent for females. Additional differences with Doiron and Riddell may be due to our exclusion of workers unemployed in the week of interest (third week of November).

[26]We use a bandwidth of 0.07 and a Gaussian weight function. Each observation is weighted by the product of the sample weight and the usual hours of work per week. These "hours–weighted" estimates put more weight on workers who supply a large number of hours to the market. Also all densities presented here integrate to one and thus do not reflect the relative weights of the

among Canadian women in 1988 are 43 percent for female jobs, 26 percent for mixed jobs, and 35 percent for male jobs. As argued in DiNardo, Fortin and Lemieux (1996), unionization leads to a more compressed wage structure. Correspondingly, the densities of women's wages in both female jobs and male jobs are much more compressed than the corresponding density in mixed jobs; these former densities in fact share *the same mode.* Figure 4 also shows that mixed jobs are the type of jobs comprising the larger proportion of women working at the lowest wages.[27] Finally, whichever the job type, the densities of women's wages are shifted towards the lower end of the distribution compared to those of men's wages.

Differences in the occupational characteristics of the jobs in which women and men work have also been investigated as a potential explanation of the effect of gender composition on wages. Women may earn less because they work in occupations which require less skills and are thus less productive or valuable to the firm (Hodson and England 1986). Men may earn more because they work in riskier jobs (Leigh 1984), that carry compensating wage differentials. To investigate this possibility we also examine the contribution of some important job characteristics from the *Canadian Classification and Dictionary of Occupations* (CCDO) (the Canadian equivalent of the *Dictionary of Occupations Titles (DOT)*). As explained in more detail in section 5.2, we extract the following characteristics from the CCDO: general educational development (GED), specific vocational preparation (SVP), physical demands, and environmental conditions. The GED and SVP were available from the Strategic Policy Group at Human Resources Development Canada in machine-readable form. The other characteristics, however, had to be typed in from the various manuals and their updates.[28] The job characteristics are available for the seven-digit occupation

types of jobs.

[27]This results hinges on whether the occupation "sales clerk" is a mixed or a female job; this may vary by province.

[28]While Hunter and Manley (1986) have made a machine-readable version of 43 CCDO worker-trait items available, their version relates to the 1971 SOC and does not include environmental

codes (more than 6,500 categories) and, in the absence of appropriate weights, have to be averaged over the four-digit categories.[29] Although the reliability of the CCDO occupational characteristics has yet to be assessed, they are likely to have the same problems (i.e., gender bias) as their DOT counterparts (see, e.g. Miller, Treiman, Cain and Ross (1980)).

4. Econometric Framework

Drawing from the different perspectives of standard human capital theory and of personnel economics (or human resource management), we include both individual and job characteristics in our model of wages. The log wages of individual i are

$$\ln w_i = X_i\beta + \alpha_k \cdot OCC_{ki} + v_i \tag{1}$$

where the X_i are characteristics which vary by individual, OCC_{ki} are occupation dummies which take the value 1 if the individual is in occupation k and 0 otherwise, and v_i is an individual specific error term. The relationship between the occupation fixed effects, α_k, and the gender composition of that occupation, which is our primary interest, is specified as

$$\alpha_k = \lambda + \gamma PFEM_k + \eta_k \tag{2}$$

where $PFEM_k$ is the percentage of workers in occupation k who are female, and η_k is an occupation wide error term. Substituting (2) into (1), we obtain

$$\ln w_i = \lambda + X_i\beta + \gamma PFEM_k + (\eta_k + v_i). \tag{3}$$

conditions.

[29]Note that a similar procedure was used in Macpherson and Hirsch (1995).

It is clear that the standard errors obtained from ordinary least-squares (OLS) estimation of this equation would be biased, as the error term is correlated across individuals within occupations due to η_k.[30]

One way to proceed would be to estimate (3) directly by generalized least–squares (GLS). An alternative is the following two–step procedure. First, estimate equation (1) by OLS, or in our case weighted least-squares (WLS) as we use the LMAS supplied individual level weights in the estimation. We can express the resulting estimates of the occupation effects as

$$\widehat{\alpha}_k = \alpha_k + \epsilon_k, \tag{4}$$

where ϵ_k is the measurement error in the $\widehat{\alpha}_k$. We then estimate the equation

$$\widehat{\alpha}_k = \lambda + \gamma PFEM_k + (\epsilon_k + \eta_k), \tag{5}$$

substituting our estimates of the occupation effects for the dependent variable in equation (2). Note that the measurement error in the dependent variable does not bias the estimate of γ. The appropriate estimation strategy for (5) depends on which error component, ϵ_k or η_k, dominates the composite error term. On the one hand, ϵ_k is likely to be heteroskedastic which would suggest a GLS strategy. In this case the appropriate weights are proportional to an occupation's sample size or the variance of its fixed effect α_k. On the other hand, there is no obvious reason why η_k should not be homoscedastic, and so if it dominates, OLS, or what we will call unweighted least squares (UWLS) for reasons which will become clear, is appropriate for the second stage. In this strategy each occupation would be weighted equally. [31]

[30]Since we would use sample weights in this regression, it would strictly speaking be a weighted least squares regression.

[31]This strategy thus takes jobs as unit of observation rather than individuals. For problems with

To provide a comparison, we present results using UWLS and two feasible GLS estimators in the second stage regressions. In GLS1 we use the WLS estimates of the sampling variances of $\widehat{\alpha}_k$ from the first stage regressions as weights.[32] In GLS2 the sum of the LMAS sample weights (by occupation) are used as weights. Note that our econometric strategy accounts for the problem of using grouped data in an individual level regression, as noted by Moulton (1986). This problem is acknowledged in Macpherson and Hirsch (1995) (p.450) who when using a two-step procedure obtain standard errors 10 times larger than for the one-step WLS estimates.[33]

5. Results

5.1. *Estimates of the PFEM Wage Penalty*

In table 3 we present the results of the second stage regressions, the estimated relationship between wages and the femaleness of employment in Canada, progressively adjusting for individual level productivity characteristics in the first stage regressions. We begin in the first row controlling for "human capital" variables: a quartic in age and six education

this type of analysis, see Cheng, Orazem, Mattila and Greig (1997). Also, note any weaknesses of the occupation classification system will carry into the estimation. The occupation classification systems used in this study are male biased in that they classify blue collar workers at a more detailed level than white collar workers. More precisely, there are 299 male occupations, 133 mixed occupations and 80 female occupations in our Canadian sample. American 3-digit occupation codes are vulnerable to similar criticism.

[32] Since the first stage regressions are estimated by weighted least-squares using the LMAS sample weights, following Wooldridge (1998) it might be preferable to use White estimates of the sampling variances of the $\widehat{\alpha}_k$ as weights in GLS1. Note, however, that many of the occupation cell sizes are very small so the finite sample bias of the White estimates could be quite severe. We have experimented with this procedure and in practice found that it yields results very similar to the UWLS estimates reported in table 3 (i.e., it weights the different occupations fairly evenly).

[33] Macpherson and Hirsch (1995) also report changes in the estimated coefficients; for example, the gender composition coefficient for males from their expanded specification goes from -0.0986 with the one-step to -0.1305 with the two-step.

classes.[34] At the aggregate level, the message of these results will turn out to be the message of our paper: the $PFEM$ penalty for females is small and statistically insignificant, while the penalty for males is much larger. Given an average $PFEM$ of 0.25, the estimates for males, which average roughly -0.23 across the two years, imply an elasticity of -0.058 (0.25×-0.23). Furthermore, we can easily reject the hypothesis that any of the estimates is equal to zero. In contrast, the three estimation strategies are in agreement that the penalty for females is effectively zero in 1988. Note, however, that the 1987 estimates provide some conflicting evidence: the UWLS result is negative and significantly different from zero. Finally, similar to the findings of U.S. studies, the largest changes in the estimate of γ with the inclusion of human capital variables, relative to the unconditional estimates in table 1, are for males.

In the second row for each year we add explanatory variables in an attempt to replicate the conditions in which a pay equity policy might be implemented. Their target is the relationship between wages and $PFEM$, net of differences in allowable productivity related characteristics. Therefore, we attempt to control for systematic variation in wages across firms and with job/individual characteristics which are likely to be tolerated in the representative legislation. Johnson and Solon (1986) show that this exercise highlights the limitations of pay equity policies. In particular, much of the correlation of wages and $PFEM$ is across industries and firms, and thus outside the purview of most legislation.

The additional explanatory variables in these regressions are province effects, 11 industry effects and dummy variables for metropolitan area, employment in the federal, provincial or local governments, union coverage and part time status. The effects of this change in specification are smaller parameter estimates for each group. The largest changes

[34]The unadjusted estimates are reported in table 1. Note in this case when there are no other conditioning variables, the estimates from GLS1 and GLS2 match the estimates we obtain simply regressing individuals' wages on $PFEM$ weighting by the LMAS individual level weights reported in table 1. The returns to the human capital variables are reported in Baker and Fortin (1999b).

are observed for males.

In the last specification we add demographic variables, some of which are unlikely to be considered legitimate bases of wage variation in legislation. These include tenure, firm size, the numbers of preschool and older children respectively (up to 3) (for 1988) and dummy variables for marital status and visible minority status. In each year, and for either gender, the effect of these new variables is very small. The estimates of γ remain essentially unchanged.

In attempting to summarize these results it is necessary to reconcile any differences across years, and in some instances across the different estimation strategies. The original discussion of the different strategies was couched in terms of efficient estimation, and thus asymptotically they should lead to the same estimates. In this light any difference in the results from the three procedures should be viewed as a finite sample phenomenon. Another possibility, however, is that they estimate different objects. The UWLS approach weights each occupation fixed effect equally, while GLS2 weights them in proportion to the (weighted) sample size of the occupation. GLS1 walks a middle ground as the the WLS estimates of the sampling variances of the $\widehat{\alpha}_k$ (from the first stage regressions) should be proportional to occupational sample size. If γ is the same across all occupations, irrespective of size, then the weighting strategy is irrelevant. If there is parameter heterogeneity, however, the UWLS procedure estimates the average wage penalty to $PFEM$ across all occupations, while the GLS2 procedure estimates the penalty faced by the average individual.

The major discrepancy in the results for males is in the estimates across years. In the richer specifications, the 1987 results are generally smaller than their 1988 counterparts; roughly by one-half using the UWLS estimation strategy. A limitation of the LMAS data is that small sample sizes mean that the same occupations are not necessarily observable in both years, and the the estimate of mean wages can change dramatically for those that are. The first problem is clearly evident here as the number of occupations drops from 473 to

456 between 1987 and 1988. This difference in occupational composition appears to play a small role in a reconciliation. There are 453 occupations that are observable in both years. Limiting the sample to these occupations and using the third specification and the UWLS estimation strategy leads to an estimate of -0.091 (0.037) for $\widehat{\gamma}$ in 1987 and -0.150 (0.037) in 1988. A second consideration is that the 1987 results are sensitive to a few observations.[35] Simply excluding four influential but small occupations leads to an estimate of γ of -0.114 (0.036) (again using UWLS and specification three). A similar analysis of the 1988 results reveals that the estimates are not so obviously influenced by a few observations, and of the four sensitive occupations identified in the 1987 data, only Dental Hygienists and Technicians turn up again as important to the 1988 result. Excluding this occupation leads to $\widehat{\gamma} = -0.140$ (0.037). It is troublesome that the estimates are sensitive to the inclusion of such small occupations, which at the same time underlines the weakness of an estimation strategy, such as UWLS, that does not account for occupational sample sizes. While excluding them is certainly arbitrary, the preceding arguments suggest that the 1988 results may serve as better summary estimates of γ for males.

The potentially more controversial reconciliation is for females. Most of the estimates suggest the wage penalty for $PFEM$ is quite small and statistically insignificant; the exception is the UWLS results for 1987. In this case the number of occupations is quite stable over the two periods, although there are changes in composition. In fact, only 331 occupations are present in both years. Again, using specification three as a basis of comparison, the UWLS estimate of γ for 1987 using the common occupations is -0.083 (0.048), and for

[35]A useful measure of the influence of an observation is the DFBETA which measure the difference between the regression coefficient, here $\widehat{\gamma}$, when the ith observation is included and excluded. This difference is then scaled by the estimated standard error of the coefficient. An examination of the DFBETA's identifies four occupations,—Audio and Speech Therapists (0.91), Dietitians and Nutritionists (0.94), Dental Hygienists and Technicians (0.97), and Inspectors, Testers, Graders and Sorters: Other Processing Occupations (0.64)—, as particularly influential on the results ($PFEM$ reported in parentheses). These influential occupations were identified by examining cases where the absolute value of the DFBETA was greater than $2/\sqrt{n}$.

1988 is -0.038 (0.053). Not surprisingly, in both years the occupations excluded in these regressions tend to be male jobs. Also, there are not particularly influential observations in either year, with the exception of Dancers and Choreographers in 1988.[36] Excluding this occupation from the 1988 sample leads to an UWLS estimate (specification three) of -0.055 (0.050). Certainly the weight of the evidence suggests that the $PFEM$ wage penalty for Canadian females, or at least the penalty faced by the average female, is modest. In fact, we cannot reject the hypothesis that it is equal to zero.

In the rest of our analysis, we focus on 1988 and only report GLS2 results, as carrying all three estimators becomes increasingly unwieldy. In general, the GLS2 estimates are representative of the inference from the different approaches for that year. Finally, in those cases where there is some sensitivity to the estimation strategy, the straightforward interpretation of the GLS2 estimates—the wage penalty for $PFEM$ faced by the average individual—is likely of greater interest from a policy perspective.

5.2. The Effects of Occupational Characteristics

One explanation for the correlation of wages and occupational gender composition is that it reflects returns to unobserved skills or compensating wage differentials for as yet excluded occupational characteristics. In fact, Macpherson and Hirsch (1995) argue that as much as one-quarter of the relationship for females and one-half the relationship for males is due to these sorts of factors. Furthermore, they argue that once control for detailed occupational characteristics is made, the relationship is generally larger for females than for males—just the opposite of the conventional wisdom.

We examine this issue in a Canadian context in table 4. In the first row (specification three) we start from the final rows of table 3 and add controls for the CCDO skill requirements characteristics: general educational development (GED), measured in approximate

[36]This conclusion was reached examining the DFBETAs.

of years of schooling, and specific vocational preparation (SVP), measured in months of training. In Canada, controlling for skill requirements decreases the magnitude of γ for females but increases it for males. Macpherson and Hirsch (1995) found these sorts of controls decreased the estimated relationship between wages and gender composition for both males and females. In specification 5, we add a control for hazards defined in terms of the CCDO sixth category of environmental conditions as situations in which the individual is exposed to the definite risk of bodily injury. This control decreases the magnitude of the $PFEM$ coefficients for males but leaves the estimate for females unchanged. Note that the result for males–the positive and significant effect of hazards on wages–is consistent with a compensating wage differentials story. In the sixth specification, we use the following controls for strength and physical demands: sedentary work–medium work, heavy work, bending, visual skills and motor coordination.[37] Finally, in specification 7 we add controls for outside and inside work, corresponding to the CCDO work location variable (EC-1). Overall, these additional controls lead to an estimate of γ for females which is essentially 0, although the estimate was small and statistically insignificant before they were added. For males the additional controls have virtually no effect on the estimated relationship between wages and occupational gender composition.

Therefore, this analysis has uncovered yet another difference between the Canadian and U.S. cases. Relative to the estimates in table 3, the ultimate contribution of the additional occupational characteristics is effectively zero for males. There are modest effects on the estimates for females, but we are comparing two statistically insignificant coefficients.

[37]Following a multifactorial analysis of the original CCDO codes we constructed the following variables. Using the CCDO codes, in the physical activities (PA) category, sedentary work-medium work corresponds to PA-1: S,S-L,S-M; heavy work to PA-1: H and VH; bending to PA-3; visual skills to PA-7; and motor coordination to the sum of PA-2-4-8.

5.3. *Heterogeneity in the Effect of Gender Composition on Wages*

An objection to the analysis thus far is that we are failing to capture any heterogeneity in the effects of gender composition on wages across groups; for example, union/nonunion or full-time/part-time differences. Furthermore, it's possible that the very small estimates of γ we obtain for females result from these sorts of differences; if we focus on full time workers we may recover the "expected" larger negative estimates.

In table 5 we present estimates of γ for males and females in 1988 by these different groupings to investigate this possibility. The first panel contains the results by age groups. For females the penalty is modestly larger for those of prime working age. For males there is evidence that it is the penalty faced by young men that drives the aggregate results.

The second panel contains the results by education. The penalty for females follow a U shaped pattern and is notably severe for those with a university degree. Interestingly, the penalty for university educated males is very similar. That said, the penalties to less educated males are also quite substantial in contrast to the results for females.

In the third panel we investigate differences by union status. For both males and females the penalty in non-union jobs is substantially greater than in union jobs. In fact, using the specification conditioning on the sectoral controls as a point of comparison, union work would appear to pay a modest premium to women in female jobs.

Finally in the last panel we present estimates by full-time status. Again using the sectoral controls specification to compare, full-time men and women in female jobs face very similar penalties. In part-time work, however, there is a premium to female jobs for women and a penalty for males.

What happens if we combine some of these splits of the data? When we restrict our attention to the sub-sample of full-time non-unionized women (47 percent of working women), we find estimates of γ ranging from -0.236 to -0.250 (with standard errors of around 0.06). If we further restrict the sample to full-time non-unionized women with a university degree (who are not particularly low wage workers and represent 11 percent of working women), we find estimates of γ ranging from -0.315 to -0.336 (with standard errors of around 0.1).

These results do reveal some heterogeneity in the penalty to female jobs across different groups of men and women, but perhaps not in ways that we might expect. Consistent with our aggregate inference, in most decompositions the penalty for males is larger than that for females. The groups of women who face penalties comparable to their counterparts in the United States (at the aggregate level) are university educated. More generally, larger penalties for females are found in full-time work and the non-union sector. Varying the sample allows us to retrieve estimates ranging from -0.336 to 0.353.

5.4. A Comparison to the Results of Other Studies

How do these results compare to those of other Canadian studies? As noted above, the Canadian literature is very small. Baker et al. (1993) obtain an estimate of γ for females of 0.055 (0.034) in 1985 using the gender composition of 2-digit occupations. This is not inconsistent with our small estimates for females, as they also show that the estimated parameter on $PFEM$ is more positive (less negative) as you aggregate occupations, and our results are based on the gender composition of 4-digit occupations.

Reilly and Wirjanto (1999) report much larger estimates of a penalty to "female work" for the Maritime provinces in 1979: -0.293 (0.110) for females and -0.223 (0.083) for males. In addition to the difference in year, differences in sample and the definition of $PFEM$ complicate direct comparison of these results with the estimates here. First, Reilly and Wirjanto's results are for three provinces. Second, they restrict their sample to full time workers. Third, they use a smaller (and different) set of control variables. Finally, and perhaps most importantly, Reilly and Wirjanto measure $PFEM$ at the establishment, rather than the occupational level. Conditional on some broad occupational controls, their definition of a female job is work in a firm that has a high proportion of females. Our definition of a female job is an occupation that has a high proportion of females, even if the work is completed at a mostly male firm. Of course, the two definitions are related as many women who work in female occupations may also work in female firms. Alternatively, the Reilly and Wirjanto study speaks to the debate surrounding policies such as affirmative action and employment equity, while the analysis here is related to the debate over comparable worth/pay equity programs.

Much more straightforward comparisons are possible to the large U.S. literature. The results that are probably most directly related are Macpherson and Hirsch's (1995). In their "standard" specification, which is roughly comparable to our specification three, the estimated $PFEM$ penalty for American females in 1987 is -0.170, and for males is -0.219. The corresponding results for 1988 are -0.168 and -0.194, respectively. Therefore, males and especially females in the United States face a much larger penalty for working in female jobs.[38]

6. Gender Gap and Gender Composition

Pay equity/comparable worth legislation has been enacted in some jurisdictions in an attempt to reduce the gender gap, understood to be mainly caused by occupational segregation. The specific target and the evaluation of these policies thus is typically debated against the background of the gender wage gap. There is some interest, therefore, in discovering how $PFEM$ contributes to the difference in wages between males and females.

From our first stage regressions we have

$$\overline{\ln w^j} = \widehat{\beta}^j \overline{X^j} + \widehat{\alpha}_k^j \cdot \overline{OCC_k^j}, \tag{6}$$

[38] In Baker and Fortin (1999b) we have estimated the U.S. wage penalties using our CPS-ORG samples, using more directly comparable specifications of the estimating equations. The results reveal similar U.S./Canada differences.

where we now add superscripts to distinguish estimates for males and females ($j = M, F$) and the overbar denotes the relevant mean. This implies

$$(\overline{\ln w^M} - \overline{\ln w^F}) = (\widehat{\beta}^M \overline{X^M} - \widehat{\beta}^F \overline{X^F}) + (\widehat{\alpha}_k^M \cdot \overline{OCC_k^M} - \widehat{\alpha}_k^F \cdot \overline{OCC_k^F}). \tag{7}$$

The second term on the right hand side of (7) is just that part of the log wage differential that is accounted for by differences in the occupation effects and the distribution of individuals across occupations. Similarly, from the second stage regressions we have

$$\overline{\widehat{\alpha}^j} = \widehat{\lambda}^j + \widehat{\gamma}^j \cdot \overline{PFEM^j}. \tag{8}$$

A standard Oaxaca decomposition of the second stage equations for males and females yields

$$(\overline{\widehat{\alpha}^M} - \overline{\widehat{\alpha}^F}) = (\widehat{\lambda}^M - \widehat{\lambda}^F) + \widehat{\gamma}^M(\overline{PFEM^M} - \overline{PFEM^F}) + \overline{PFEM^F}(\widehat{\gamma}^M - \widehat{\gamma}^F). \tag{9}$$

Equations (7) and (9) are related by noting that $\widehat{\alpha}_k^j \cdot \overline{OCC_k^j}$ in (7) is implicitly the sum $\sum_{l=1}^K \widehat{\alpha}_l^j \cdot \overline{OCC_l^j}$, and that $\overline{\widehat{\alpha}^j} = \sum_{l=1}^K \widehat{\alpha}_l^j \cdot \overline{OCC_l^j}$ when we use GLS2 to estimate the second stage regression.[39] Therefore, under the GLS2 weighting scheme equation (9) provides a decomposition of that part of the log wage gap that is accounted by male/female differences in both occupational employment and occupational returns. Note also from (8) that

$$(\widehat{\gamma}^M \cdot \overline{PFEM^M} - \widehat{\gamma}^F \cdot \overline{PFEM^F}), \tag{10}$$

is just that part of the wage gap due to differences in both the average femaleness of employment and the associated penalties.

[39]Note that OCC_k^j is a 0/1 variable.

One way of viewing (10) is as an (ceteris paribus) estimate of the potential effect of policies aimed at eliminating the correlation of wages with $PFEM$ on the log wage differential (i.e. if $\gamma^M = \gamma^F = 0$).[40] Estimates of (10) are easily constructed for 1988 using average $PFEM$ from table 1 and the GLS2 estimates of γ^j for this year from table 3. The estimates range from -0.04 to 0.02.[41] Therefore, the aggregate effect of γ and $PFEM$ is very modest. As can be seen in tables 1 and 3, while females are penalized by a much larger average value of $PFEM$, they gain from having much smaller estimates of γ. Since the log wage gap in Canada was 0.27 in 1988, these results suggest that eliminating the effects of gender composition on wages would have limited effect on the log wage differential.

Following previous studies, in table 6 we present the Oaxaca decomposition's represented by (9). Here we isolate that part of the wage gap that can be associated with differences in $PFEM$ across the genders. The policy implications of these results are less clear. While employment equity programs have a stated objective of increasing the representation of females in certain occupations it seems unlikely that the end result would be $\overline{PFEM^M} = \overline{PFEM^F}$. Macpherson and Hirsch (1995) report that differences in $PFEM$ account for roughly 0.08 log points of the U.S. log wage gap in 1988. Our estimates are generally smaller, except in the "Human Capital" specification. This is due, in part, to the fact that we weight the difference in $PFEM$ by $\hat{\gamma}^M$, and that the GLS2 estimates of this parameter (table 3) are smaller than both Macpherson and Hirsch's result and the GLS1 estimates [42] In Canada, differences in $PFEM$ account for between 0.04 to 0.09 log points of the gender log wage gap. Note that in specifications 2 and 3 the aggregate impact of the occupation effects and the distribution of females across occupations increases the wages of females relative to males.

[40]Note we are ignoring any obstacles pay equity policies might face in achieving this goal. See, for example, Johnson and Solon (1986).

[41]The estimates are -0.042, 0.019 and 0.014 for specifications one through three respectively.

[42]Note that Macpherson and Hirsch (1995) use a weighted average of the male and female estimates.

7. Discussion and Conclusions

Our analysis of the effect of gender composition on wages in Canada has uncovered some surprising differences from the evidence for other countries. Although there is some heterogeneity across subgroups, most of the estimates of the relationship between wages and gender composition for Canadian women are quite small and typically statistically insignificant. On the other hand, we find uniformly negative and substantial penalties for males in female jobs.

These results contrast with estimates from the large U.S. literature in this area. For example, Macpherson and Hirsch (1995) report that males and females in the United States faced quite similar, negative effects of gender composition on wages in 1987 and 1988. While our estimates for males are roughly comparable to this evidence, the estimates for females are quite different. In Baker and Fortin (1999c) we argue that a large part of the Canada/U.S. difference for females is accounted for by cross-country differences in unionization and differences in the wage premiums to certain "public goods" sector jobs. Macpherson and Hirsch also report substantial reductions in the wage penalty to female jobs when they control for detailed job characteristics, suggesting the lower wages in female jobs are related to (typically) unobserved job attributes or requirements. In contrast, our evidence for Canada reveals the addition of detailed occupational characteristics has quite modest effects on our inference.

Analysis at a finer levels of aggregation does reveal some heterogeneity in the penalty across groups. Most notably, females who are university graduates face relatively large and significant penalties to working in female jobs. Larger, negative wage effects are also found for women in the non-union sector, and those who work full time. The penalty for males is more uniform across different decompositions of the data.

Our evidence is from a period (1987-1988) when the labour market was mostly untouched by the effects of pay equity legislation. Since this time, pay equity programmes have been introduced to the public sectors of most provinces, the private sector of Ontario, and more recently to the private sector of Quebec. The results of this study provide some perspective on the "target" for these initiatives, as well as any further extensions of comparable worth in other provinces. The nominal purpose of this legislation is the elimination of the negative wage penalty in female jobs. For Canadian women this is a small target. Simulations of the effect of eliminating the relationship between wages and gender correlation on the gender wage gap correspondingly suggest little effect on the relative economic stature of women. Although pay equity legislation typically does not distinguish the sex of individuals working in female job classes, the more popular view of this policy is that it will aid low wage women. Surprisingly, a larger target for these policies may be the wage penalty Canadian males experience in female jobs. Similar counter-intuitive suggestions flow from the inference at the disaggregate level. Some women do face substantial penalties to working in female jobs. An outstanding example is the university educated. These are not the individuals that are typically thought of as toiling in underpaid female jobs.

Our results do not preclude the existence of low paid female jobs, such as waitresses and cashiers. However since low paid female jobs have counterparts in mixed and males jobs for which women are equally low paid, such as sales clerk (mixed) and material handling jobs (male), and since there are female jobs that are relatively higher paid, such as nursing and teaching jobs, the gender composition of employment does not have a strong consequence for low pay for females at the aggregate level. Studies documenting the contribution of inter-firm wage differentials to the gender wage gap suggest that pay equity/comparable worth legislation may be a relatively ineffective remedy.[43] Our results identify further obstacles to these policies in Canada.[44]

[43]See Reilly and Wirjanto (1999) for Canada, and Carrington and Troske (1995) and Petersen and Morgan (1995) for the United States.

[44]Baker and Fortin (1999a) provide an analysis of the effects of the introduction of pay equity to the private sector in Ontario.

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Table 1

Canada/U.S. Comparison of Mean Wages, Gender Composition, Wage–Composition Relationship and Wage Gap by Job Types

	Women					Men					
Sample	N	Wage	PFEM	$\hat{\gamma}$		N	Wage	PFEM	$\hat{\gamma}$		Female/ Male Wage Ratio
CANADA:1987											
All jobs	17810	9.87	.676	.006	(.061)	21500	13.03	.254	-.130	(.052)	.758
Female jobs	10801	9.95	.858	-.006	(.337)	1627	12.36	.773	-.342	(.427)	.805
Mixed jobs	5617	9.47	.467	-.792	(.369)	6277	13.04	.437	-.492	(.359)	.726
Male jobs	1392	10.76	.190	.758	(.251)	13596	13.11	.091	.110	(.151)	.821
CANADA:1988											
All jobs	14868	10.88	.668	-.028	(.060)	17739	14.23	.251	-.145	(.052)	.765
Female jobs	8815	10.91	.857	-.082	(.320)	1324	13.94	.777	-.603	(.399)	.783
Mixed jobs	4876	8.72	.465	-.992	(.381)	4963	13.89	.435	-.780	(.364)	.764
Male jobs	1177	9.69	.189	.913	(.156)	11452	14.41	.099	.175	(.156)	.818
UNITED STATES:1987											
All jobs	80009	9.70	.675	-.228	(.062)	87713	13.55	.265	-.022	(.069)	.716
Female jobs	50877	9.07	.841	.175	(.271)	7899	11.76	.742	-.844	(.315)	.771
Mixed jobs	22875	10.90	.438	-.065	(.318)	29615	15.14	.405	-.199	(.377)	.719
Male jobs	6257	10.53	.191	-.501	(.295)	50199	12.90	.108	-.130	(.228)	.816
UNITED STATES:1988											
All jobs	76979	10.16	.670	-.227	(.062)	84009	14.01	.266	-.028	(.069)	.725
Female jobs	48518	9.52	.839	.130	(.278)	7498	12.00	.743	-.812	(.337)	.793
Mixed jobs	22311	11.33	.436	-.059	(.310)	28341	15.69	.404	-.205	(.381)	.722
Male jobs	6150	10.93	.187	-.292	(.288)	48170	13.35	.108	-.093	(.231)	.818

Note: Average wages in 1988 Canadian dollars (exchange rate used is 1.2174). Calculations are from the 1987 and 1988 LMAS for Canada and from the 1987 and 1988 CPS ORG for the United States. The estimated γ from the OLS and feasible GLS are identical. The corresponding estimated standard errors, in parentheses, are from the two stage estimation strategy that used the sum of the individual level (i.e., LMAS or CPS) weights (by occupation) as weights.

TABLE 2

MEANS OF SELECTED VARIABLES

Variable	Women 1987	Women 1988	Men 1987	Men 1988
Wage (1988 Canadian$)	10.32	10.88	13.57	14.23
Age	36.4	36.5	37.2	37.2
Education:				
Primary	.070	.063	.108	.104
Some High School	.104	.101	.142	.130
High School Grad	.381	.362	.348	.341
Some Post-Secondary	.107	.101	.096	.097
Post-Secondary Degree	.189	.210	.143	.162
University Degree	.149	.164	.164	.167
Part-time	.217	.226	.041	.042
Married	.670	.665	.710	.690
Visible Minority	.054	.052	.052	.051
Metropolitain Area	.652	.731	.630	.703
Industrial Sector:				
Agriculture, Forestry and Fisheries	.013	.011	.025	.023
Mining	.006	.006	.029	.029
Construction	.013	.017	.086	.085
Manufacturing				
Nondurable	.080	.073	.102	.110
Durable	.049	.047	.156	.159
Transportation and public utilities	.050	.046	.111	.116
Trade	.163	.161	.156	.156
FIRE	.088	.088	.042	.040
Business and professional services	.056	.062	.042	.043
Consumer services	.117	.121	.061	.055
Medical, welfare, and educational services	.294	.291	.097	.098
Public administration	.070	.075	.093	.086
Federal	.020	.020	.041	.042
Provincial (State)	.028	.029	.029	.023
Local	.016	.016	.034	.035
Union coverage	.361	.371	.441	.452
Tenure	5.67	5.78	7.67	8.00
Establishment Size:				
$s < 20$	.379	.376	.321	.300
$20 <= s < 100$	.316	.298	.330	.320
$100 <= s < 500$	.206	.203	.231	.237
$s >= 500$	.099	.122	.118	.142
No. of observations	17,810	14,868	21,501	17,739

TABLE 3
ESTIMATED EFFECTS OF GENDER COMPOSITION
ON WAGE LEVELS IN CANADA

	Specification:	UWLS	GLS1	GLS2
1987: WOMEN				
1:	Human capital	-.146	-.091	-.004
		(.057)	(.052)	(.047)
2:	1+ Sectoral Controls	-.108	-.056	-.040
		(.051)	(.045)	(.036)
3:	2+Individual characteristics	-.120	-.066	-.041
		(.049)	(.043)	(.034)
	No. of occupations		380	
1988: WOMEN				
1:	Human capital	-.013	-.013	-.023
		(.060)	(.055)	(.046)
2:	1+ Sectoral Controls	-.037	-.012	-.066
		(.054)	(.050)	(.037)
3:	2+Individual characteristics	-.033	-.012	-.062
		(.051)	(.047)	(.035)
	No. of occupations		378	
1987: MEN				
1:	Human capital	-.207	-.229	-.217
		(.042)	(.040)	(.036)
2:	1+ Sectoral Controls	-.081	-.099	-.052
		(.039)	(.031)	(.033)
3:	2+Individual characteristics	-.076	-.095	-.067
		(.037)	(.034)	(.030)
	No. of occupations		473	
1988: MEN				
1:	Human capital	-.274	-.252	-.228
		(.042)	(.040)	(.038)
2:	1+ Sectoral Controls	-.159	-.141	-.100
		(.039)	(.037)	(.034)
3:	2+Individual characteristics	-.151	-.131	-.110
		(.037)	(.035)	(.031)
	No. of occupations		456	

Note: Estimated standard errors are in parentheses. UWLS and GLS refer to the estimation strategy used in the second stage regressions. For GLS1, the observations are weighted by the OLS estimates of the sampling variances of the dependent variable from the first stage regressions. In GLS2 the sum of the individual level (i.e., LMAS) weights (by occupation) are used as weights. All the underlying first stage regressions are estimated by weighted least-squares using LMAS sample weights. Human capital conditions on a quartic in age and on six education classes. Sectoral controls add dummies for province (10), metropolitan area, industry(12), employment in the federal, provincial, and local public service, union status and part time work. Individual characteristics include dummy for married, visible minority, tenure, firm size (4), number of preschool children (up to 3), number of older children (up to 3).

TABLE 4

THE ROLE OF CCDO OCCUPATIONAL CHARACTERISTICS IN THE EFFECT OF GENDER COMPOSITION ON WAGES IN CANADA – 1988

		Women	Men
4:	3+Educational requirements[a]	-.011	-.177
		(.026)	(.025)
5:	4+Hazards[b]	.019	-.125
		(.028)	(.032)
6:	5+Strength physical demands[c]	-.036	-.155
		(.028)	(.030)
7:	6+Outside–Inside work[d]	-.025	-.118
		(.032)	(.034)
	No. of occupations	378	456

Note: The estimates presented are from the feasible GLS strategy where the sum of the individual level (i.e., LMAS or CPS) weights (by occupation) are used as weights in the second stage (ie. GLS2). Estimated standard errors are in parentheses.

[a] Educational requirements include CCDO general educational development (GED), measured in years of education and specific vocational training (SVP), measured in months.

[b] Hazards is CCDO–EC 6.

[c] Strength and physical demands include the CCDO following physical demands (PA) codes: sedentary work-medium work PA-1: S,S-L,S-M, heavy work to PA-1: H and VH; bending to PA-3; visual skills to PA-7; and motor coordination to the sum of PA-2-4-8.

[d] Outside and inside work are the CCDO–EC 1 and denote work location.

Table 5
Gender Composition Coefficients on Wages among Alternative Worker Groups – 1988

Specification: Group	NC	No controls	(1) Human Capital	(2) 1+ Sectoral Controls	NC	No controls	(1) Human Capital	(2) 1+ Sectoral Controls
			Women:				Men:	
Age:								
16–29	307	-.075	-.045	-.057	379	-.323	-.327	-.192
		(.061)	(.049)	(.041)		(.052)	(.044)	(.036)
30–44	307	-.059	-.050	-.109	417	-.037	-.149	-.055
		(.071)	(.058)	(.047)		.051)	(.043)	(.042)
44–69	246	.102	.073	.009	392	-.084	-.216	-.072
		(.079)	(.064)	(.055)		(.067)	(.056)	(.053)
Education:								
Drop-out	230	-.113	-.114	-.087	354	-.518	-.485	-.257
		(.060)	(.059)	(.052)		(.054)	(.050)	(.043)
High School	294	-.028	-.018	-.032	393	-.322	-.317	-.094
		(.052)	(.048)	(.038)		(055)	(.044)	(.040)
Post-secondary	260	.045	.045	-.001	364	-.294	-.228	-.165
		(.063)	(.058)	(.049)		(.054)	(.043)	(.040)
University	179	-.095	-.120	-.184	252	-.184	-.173	-.177
		(.081)	(.075)	(.066)		(.084)	(.078)	(.075)
Union coverage status:								
Nonunion	342	-.182	-.142	-.136	414	-.211	-.287	-.214
		(.059)	(.048)	(.042)		(.071)	(.051)	(.046)
Union	287	.025	.010	.044	415	-.016	-.097	.018
		(.061)	(.047)	(.060)		(.038)	(.030)	(.030)
Hours status:								
Part-time	211	.353	.323	.169	180	-.168	-.188	-.148
		(.099)	(.083)	(.066)		(.122)	(.108)	(.096)
Full-time	373	-.097	-.082	-.107	452	-.117	-.212	-.094
		(.058)	(.043)	(.035)		(.051)	(.037)	(.034)

Note: The estimates presented are from the feasible GLS strategy where the sum of the individual level LMAS weights (by occupation) are used as weights in the second stage (i.e., GLS2). Estimated standard errors are in parentheses. *NC* is the number of occupations.

TABLE 6

GENDER GAP DECOMPOSITIONS – 1988

Specification	
Total log wage gap	.273
0: No Controls	
Total due to Occupation Effects	.273
	(.019)
Part due to $\Delta PFEM$	.061
	(.022)
Part due to $\Delta\lambda$ and $\Delta\gamma$	.213
	(.019)
1: Human Capital	
Total due to Occupation Effects	-.416
	(.015)
Part due to $\Delta PFEM$	.095
	(.016)
Part due to $\Delta\lambda$ and $\Delta\gamma$	-.511
	(.021)
2: 1 + Sectoral Controls	
Total due to Occupation Effects	-.356
	(.012)
Part due to $\Delta PFEM$	.044
	(.014)
Part due to $\Delta\lambda$ and $\Delta\gamma$	-.400
	(.019)

Note: Standard errors in parentheses. The reported statistics are from decompositions of the GLS2 estimates of the second stage regressions (see equations (12) and (14) in the text). The specifications follow the conventions of table 3.

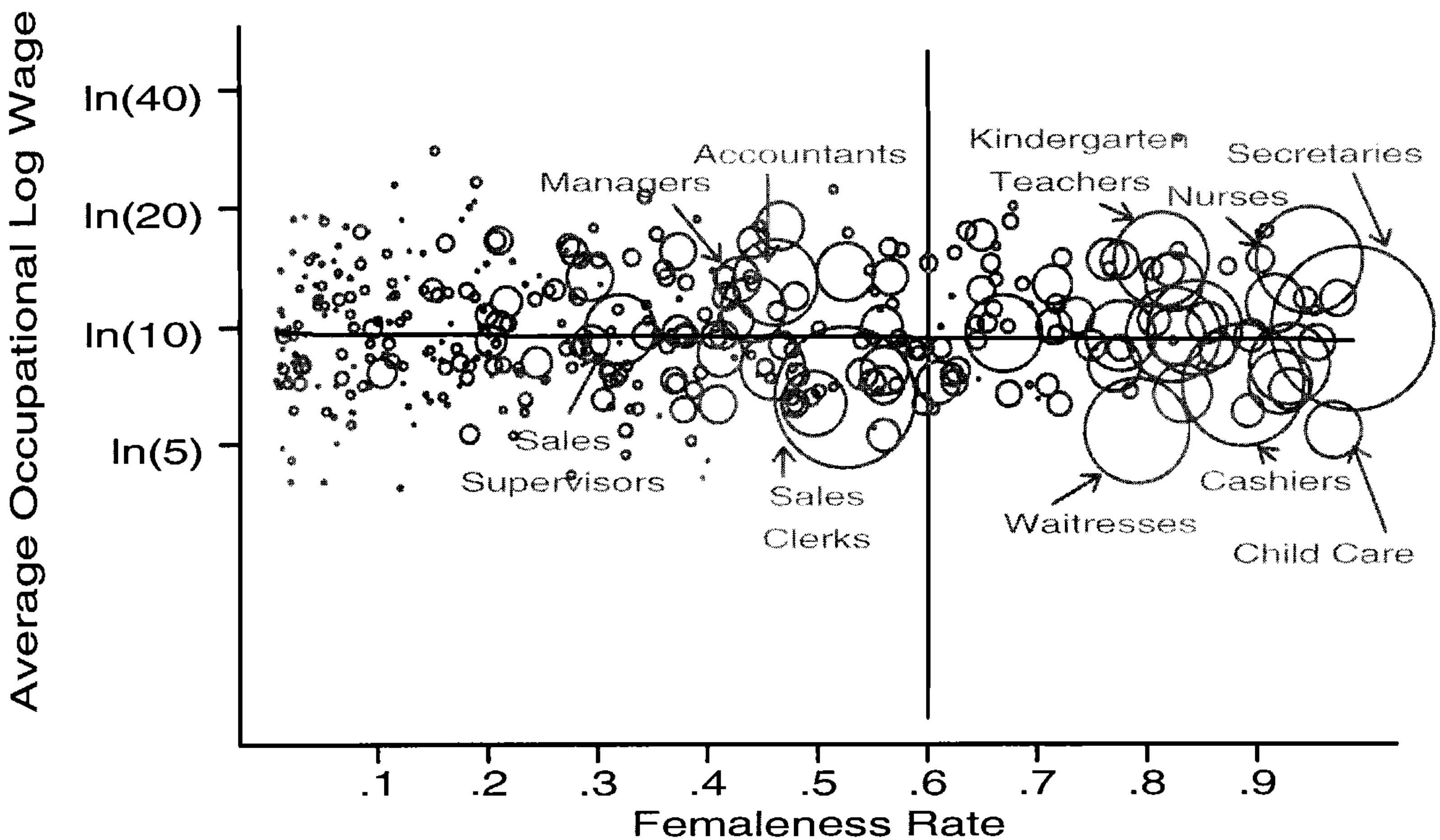

Figure 1. Effect of Gender Composition on Average Occupational Female Wages in Canada - 1988

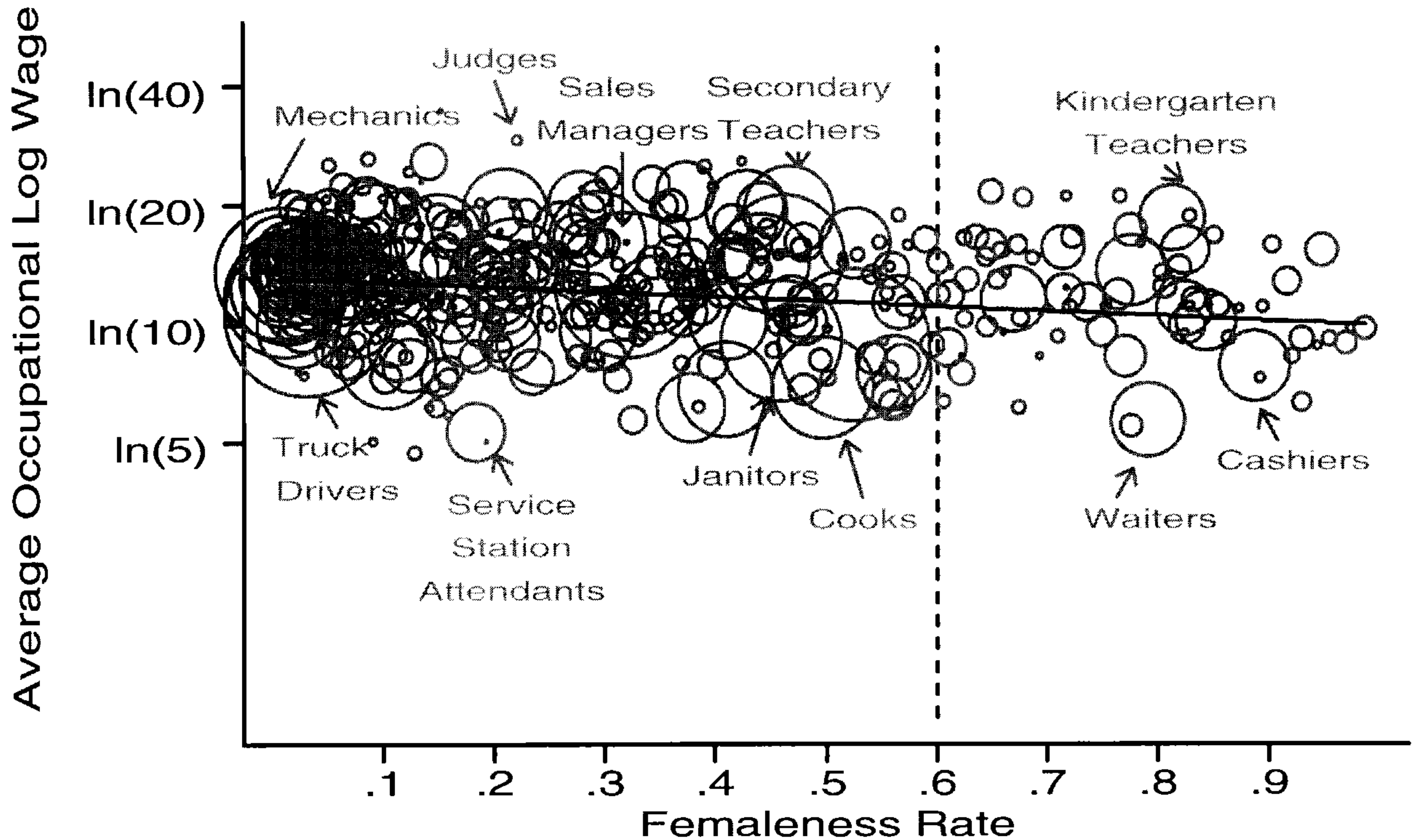

Figure 2. Effect of Gender Composition on Average Occupational Male Wages in Canada - 1988

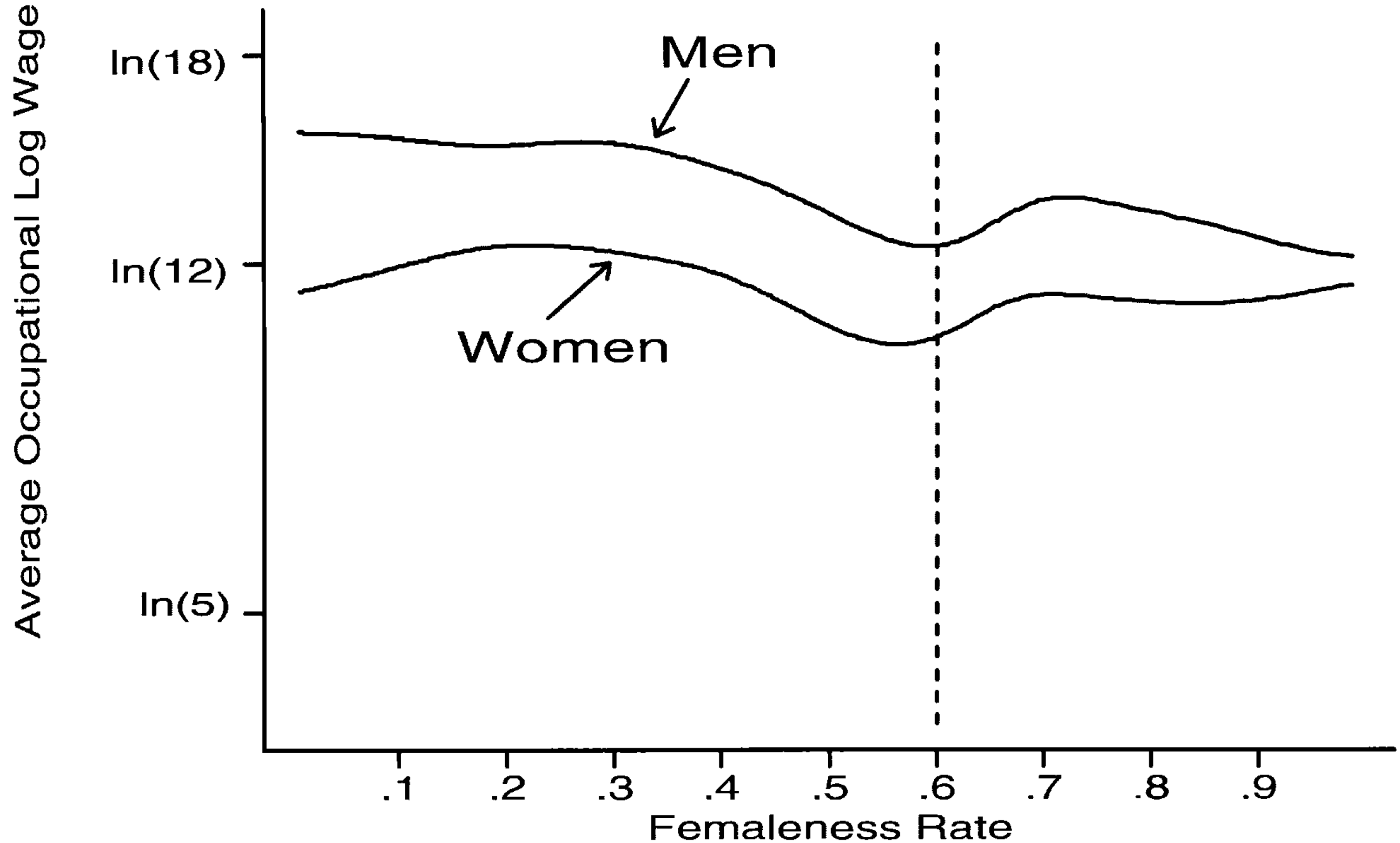

Figure 3. Weighted Kernel Regressions of Average Occupational Wages on Gender Composition in Canada - 1988

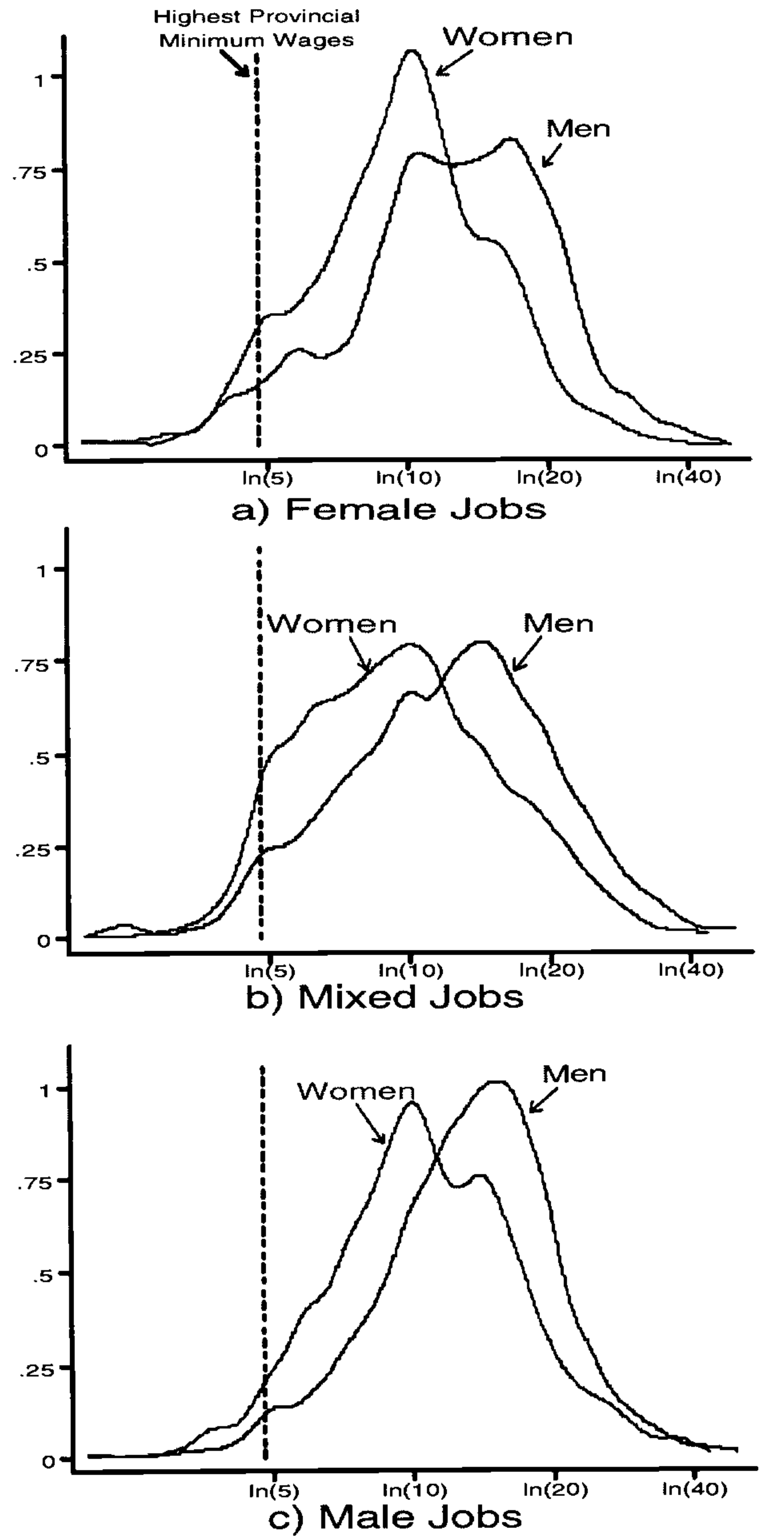

Figure 4. Kernel Density Estimates
of Log Hourly Wages - 1988
by Job Type