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Gender Composition and Wages: Why is Canada Different from the United States?

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by
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Abstract

The correlation of occupational gender composition and wages is the basis of pay equity/comparable worth legislation. A number of previous studies have examined this correlation in United States 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, the authors provide a comprehensive picture, circa the late 1980's, of the occupational gender segregation in Canada and its consequences for wages. The paper also draws explicit comparisons of canadian findings to evidence for the United States. These results indicate a link between female wages and gender composition that is much stronger in the United States than in Canada, where the relationship is generally small and not statistically significant. The relatively more advantageous position of women in female jobs in Canada is found to be linked to higher unionization rates and the industry-wage effects of "public goods" sectors.

Résumé

La corrélation entre le taux de féminité des professions et les salaires constitue le fondement de la législation en matière d'équité salariale. Diverses études antérieures ont analysé cette corrélation à partir de données américaines et ont cerné certains des facteurs déterminants des niveaux de salaire peu élevés dans les «emplois à prédominance féminine» de même que d'importantes limites des politiques publiques dans ce domaine. Toutefois, il existe peu de données visant d'autres secteurs de compétence. Cette lacune est particulièrement perturbante dans le cas du Canada, qui s'est doté de l'une des législations les plus élaborées en matière d'équité salariale au monde. Dans le présent article, les auteurs dressent un portrait complet, de la fin des années 1980, de la ségrégation professionnelle fondée sur le sexe au Canada et de ses répercussions sur les salaires. Le présent document établit également des comparaisons explicites entre les résultats pour le Canada et des données américaines. Ces résultats indiquent un lien entre la rémunération des femmes et le taux de féminité des professions beaucoup plus fort aux États-Unis qu'au Canada, où cette relation est généralement peu marquée et non significative sur le plan statistique. La position relativement meilleure dont jouissent les femmes dans les emplois à prédominance féminine au Canada est associée à des taux de syndicalisation plus élevés et aux effets fixes des salaires des secteurs des «services publics».

Acknowledgements

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We gratefully acknowledge financial support of the Centre for Interuniversity Research and Analysis on Organizations, Human Resources Development Canada, the Social Sciences and Humanities Research Council of Canada, Quebec's *Fonds pour la formation de chercheurs et l'Aide à la recherche* and the University of Toronto General Research Grant.

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1. Introduction

The casual observation that some "female jobs," such as child care work, are poorly paid is often viewed as evidence that women are "crowded" into lower-paying jobs. This belief has found more formal support in U.S. studies that document the negative effect of the "femaleness" of an occupation on wages (O'Neill (1983); Johnson and Solon (1986); Macpherson and Hirsch (1995)). As a consequence, occupational segregation has become a leading explanation of the persistence of the gender wage gap.[1] It has also engendered a policy response: comparable worth/pay equity legislation. While comparable worth programs have spread to many industrialized countries, the majority of empirical evidence, both of their curative effects and the magnitude of the problem they address, is from U.S. data. One might speculate from this development that the United States is the vanguard of legislation in this area. In fact, perhaps just the opposite is true. Canada provides a good case in point. Pay equity has been adopted throughout the public sector, and recently pro-active policies were extended to the private sectors in the provinces of Ontario and Quebec. Not only is there little evaluation of the effects of these policies, but there is, to our knowledge, no evidence that female jobs are systematically poorly paid in Canada.[2] The basis of the legislative initiatives, therefore, would appear to be the experiences of other countries.

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. We examine not only the conventional correlations between the femaleness of occupations and wage rates, but also alternative representations of the relative positions of female jobs, such as kernel density estimates. We also draw explicit comparisons of our findings to evidence from the United States. This cross country comparison helps identify the contributions of important labour market institutions, such as unions, to the correlation of the occupational gender composition with wages.

[1] Other explanations are differences between men's and women's human capital and productivity, the impact of industrial structure, and discrimination.

[2] Baker, Benjamin, Desaulniers and Grant (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), the only other study that uses detailed occupations categories, that we are aware of, finds a very small effect of percentage female on average female earnings.

We begin in Section 2 surveying the legislative environment in the two countries 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 for both Canada and the United States. They reveal that the link between female wages and gender composition is much stronger in the United States than in Canada, where it is generally small and not statistically significant. These Canada-U.S. differences are investigated in Section 6. In Section 7 we examine the relationship between the "wage penalties" in female jobs and the gender gap. We conclude in Section 8 by summarizing the Canada-U.S. differences in the effect of occupational segregation on wages and its possible causes.

2. The Legislative Environment

The objective of comparable worth 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.[3] 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 in female jobs in Canada, and any differences in these levels from their U.S. counterparts.

Canada has been called a world leader in comparable worth (e.g., Weiner and Gunderson (1990)).[4] 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 standing policies 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 the onus is placed on employers. Here there is a requirement that employers erect a pay equity plan which typically involves four steps: 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. Note that most pay equity legislation does not address wage

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[3] 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.

[4] 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).

differences across employers/establishments and industries, a potentially important source of gender wage differentials.[5]

The early complaint-based 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).[6] It is also complaint based, however, and again appears 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)).[7]

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 pay adjustments 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 in 1987 and in Nova Scotia and Prince Edward Island in 1988.[8] The implementation plans for this legislation suggest that their effects are likely outside our sample period.[9]

[5] See Reily and Wirjanto (1995) for Canada, and Carrington and Troske (1995) and Petersen and Morgan (1995) for the United States. By contrast, the pro-active Ontario legislation of 1987 allows proxy comparisons across different employers and establishments, at least in the public sector, if comparisons within the establishment are not possible.

[6] These also include crown corporations.

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

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

[9] Investigating separately the years 1987 and 1988 would permit us to see the effects, if any, of legislation passed in 1988.

Therefore, in the late 1980's Canada's labour market might be considered largely free of any effects of comparable worth policies, save for the rarely used federal and Quebec laws, and any initial effects of Manitoba's legislation.[10] It is also important to note that our sample period precedes the implementation of pro-active pay equity in the private sector in Ontario, and more recently in Quebec. 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 fully implemented until 2000.

How does this compare to the environment in the United States? There are two dimensions to be considered. First is the interpretation and application of federal laws, especially the *Civil Rights Act* and *Fair Labour Standards Act*, by the U.S. Supreme Court. The court decisions handed down throughout the 1980's are widely viewed as rejecting the principle that the federal acts encompass comparable worth. The second is the activities of state and local governments. Here the story is somewhat different. By 1987, 36 states had set up a comparable worth task force or commission, and 20 states had made some sort of pay equity awards in their public sectors (Weiner and Gunderson 1990). Thus it would appear that in contrast to current comparisons, at the time of our study, the United States was marginally ahead of Canada in pay equity policies. Certainly it is possible that public sector employment in some states as of 1987/88 would reflect the impact of comparable worth initiatives.

[10] 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 the femaleness of employment on wages in different jurisdictions over the 1980's, we can examine any provincial heterogeneity in the 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 femaleness of occupations 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).

3. Data and Descriptive Evidence

The data for this study are drawn from the Canadian Labour Market Activity Survey (LMAS) and from the U.S. Current Population Survey Outgoing Rotation Groups (CPS-ORG) for 1987 and 1988.[11] 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.[12] As explained below, additional variables measuring gender composition are obtained from Census data 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.[13] 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. Wage rates are defined similarly in the U.S. data.[14] In the U.S. data, we delete workers who had either an industry or occupation code imputed by the Census (1.3%), but we do not delete workers with imputed wages (14%) since these observations are not identified in the Canadian data.[15] The resulting sample sizes are given in table 1, which also provides the average wage levels in 1988 U.S. dollars by gender.[16] An exchange rate of 1.2174 corresponding to the spot rate of November 1988 was used.[17]

[11] Because of the rotation group format of the CPS, the 1987 and 1988 samples will be made up of the same individuals to some extent.

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

[13] That particular choice of week was dictated by comparability with other surveys in the context of a larger research project. Using the U.S. CPS-ORG, we conducted experiments to investigate potential seasonality effects. Weighted least-squares (using CPS-ORG sample weights) regressions of log wages on *PFEM* using data from different quarters leads to the following parameter estimates: -0.228 (-0.027) in Winter, -0.239 (-0.027) in the Spring, -0.230 (-0.041) in Summer, -0.212 (-0.019) in the Fall for females (and males). It would thus appear that any seasonality effect of our choice of week would be small, but admittedly a downward bias.

[14] To compute the wages of weekly earnings top coded at $999 current dollars we use unedited earnings.

[15] The LMAS data are collected through phone interviews and thus have a much lower level of allocated wages.

[16] Both the LMAS and the CPS-ORG provide sample weights that are used in the analysis.

[17] The corresponding CANSIM series label is B40001. We note that the monthly exchange rate fluctuated between 1.2853 and 1.1960 that year.

Table 1

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

Sample	Women *N*	Women Wage	Women *PFEM*	Women $\hat{g}$		Men *N*	Men Wage	Men *PFEM*	Men $\hat{g}$		Female Male Wage Ratio
Canada 1987											
All jobs	17810	8.11	.676	.006	(0.61)	21500	10.70	.254	-.130	(.052)	.758
Female jobs	10801	8.17	.858	-.006	(.337)	1627	10.15	.773	-.342	(.427)	.805
Mixed jobs	5617	7.78	.467	-.792	(.369)	6277	10.71	.437	-.492	(.359)	.726
Male jobs	1392	8.84	.190	.758	(.251)	13596	10.77	.091	.110	(.151)	.821
Canada 1987											
All jobs	14868	8.94	.668	-.028	(.060)	17739	11.69	.251	-.145	(.052)	.765
Female jobs	8815	8.96	.857	-.082	(.320)	1324	11.45	.777	-.603	(.399)	.783
Mixed jobs	4876	8.72	.465	-.992	(.381)	4963	11.41	.435	-.780	(.364)	.764
Male jobs	1177	9.69	.189	.913	(.156)	11452	11.84	.099	.175	(.156)	.818
United States 1988											
All jobs	80009	7.97	.675	-.228	(.062)	87713	11.13	.265	-.022	(.069)	.716
Female jobs	50877	7.45	.841	.175	(.271)	7899	9.66	.742	-.844	(.315)	.771
Mixed jobs	22875	8.95	.438	-.065	(.318)	29615	12.44	.405	-.199	(.377)	.719
Male jobs	6257	8.65	.191	-.501	(.295)	50199	10.60	.108	-.130	(.228)	.816
United States 1988											
All jobs	76979	8.35	.670	-.227	(.062)	84009	11.51	.266	-.028	(.069)	.725
Female jobs	48518	7.82	.839	.130	(.278)	7498	9.86	.743	-.812	(.337)	.793
Mixed jobs	22311	9.31	.436	-.059	(.310)	28341	12.89	.404	-.205	(.381)	.722
Male jobs	6150	8.98	.187	-.292	(.288)	48170	10.97	.108	-.093	(.231)	.818

Note: Average wages in 1988 U.S. 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 g 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.

We measure the femaleness an occupation (*PFEM*) as the proportion of its employment that is

PFEM

U.S. censuses (the reference years are 1990 and 1989 respectively).[18] In each case, we sample individuals who are employed in the reference week and otherwise satisfy the same selection criteria as for the job data.[19] The Canadian and American detailed occupational classifications are roughly the same order of aggregation, comprising approximately 500 categories; they are the 3-digit occupation codes in the U.S. data and the 4-digit occupation codes for Canada.[20] There are, however, notable differences in the coding of occupations across the two countries that could potentially be a factor in our analysis. For example, post-secondary teachers are classified by field in the United States while they make up only one category in Canada; blue-collar workers in Canada are classified by industry while they are not in the United States. To investigate the impact of these different classification systems, for each country we present results using both the relevant country specific occupation codes, and a "crosswalk" in which the codes for the two countries are mapped into common categories. Because of differences in the country specific codings in some instances the "crosswalk" aggregates more than one of the original categories reducing the total number of categories to a maximum of 310. Generally, this aggregation takes place across occupations with similar gender composition, but there are exceptions. For example, barbers and hairdressers, or tailors and dressmakers, that are distinct categories in the U.S. coding are aggregated into single categories in the Canadian and crosswalk coding.

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. Baker et al. (1983) provide some evidence of the relationship between 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, they demonstrate that estimates of the correlation are sensitive to the aggregation of the occupational

[18] The Canadian 1980 SOC occupational codes available from the LMAS are also available in the 1991 census. On the other hand, the 1990 U.S. Census uses the 1990 codes while the 1987 and 1988 CPS-ORG use the 1980 codes. There were fortunately only six occupational changes, which we were able to recode.

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

[20] The more detailed seven digit occupation classification system, comprising around 6,500 categories, have not been coded in any general survey that we are aware of.

categories.[21] We were fortunate to gain access to versions of the census and LMAS files that include the more detailed occupation codes.[22]

In table 1 we provide an overview of the gender composition of occupations and its consequences for wages in Canada and the United States in 1987 and 1988. Across all jobs, the femaleness rate, *PFEM*, by gender, is very similar in the two countries. For women, employment is about 67 percent female on average, while for men it is 25 or 26 percent female. The statistics are also reported by "female," "mixed" and "male" jobs. Predominantly female jobs are defined as those with a femaleness rate of 60 percent or higher.[23] 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 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.[24]

We also report average wages (in 1988 U.S. dollars) and $\hat{g}$ from the regression $\ln w_i = d + g\ PFEM_i + \epsilon_i$ estimated by weighted least-squares, using LMAS and CPS-ORG sample weights respectively. None of the differences in average wages across job types would be statistically significant given the large standard deviations, but these descriptive statistics give

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[21] 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 positive and equal to 0.354 (0.028) and 0.055 (0.034) for the 1-digit and 2-digit codes respectively (standard errors in parentheses). Similar changes are reported for the results for males.

[22] In addition to detailed occupation codes, our Canadian data also contain a single year age variable (as in U.S. data) instead of the usual 5-year classes available in the LMAS.

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

[24] The Duncan index of segregation, measured by $1/2 \sum \left| m_j - f_j \right|$, where m_j and f_j are the proportion of male and female employment, respectively, in occupation *j*, 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.

a flavour of the results to come. 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. It is thus not surprising that, for women, the estimate of g is effectively 0 in Canada, while in the U.S. the implied elasticity at an average percentage female of 0.67 is (0.67 $\times$ -0.227) -0.152. For men the two countries trade places: now in the U.S. the estimate of g is roughly 0, while in Canada the implied elasticity at an average percentage female of 0.25 is (0.25 $\times$ -0.135) -0.033. Note that the U.S. results are similar to those reported in Macpherson and Hirsch (1995) for these years.

The occupations "driving" the simple regression coefficients are illustrated in figures 1 and 2, where we plot the regression line of average occupational log wages on the femaleness rate for Canada and the United States in 1988. The relative size of the circles indicates the relative weights of the occupations. These pictures clearly show a negatively sloped regression line in the United States, while the corresponding line in Canada is flat. Note that cashiers, waitresses and child care workers all appear relatively higher paid in Canada, indicating a potential role of the minimum wage in raising the wages of the lowest paid workers.[25]

In figure 3 we plot kernel regressions of the same relation for both Canada and the United States.[26] Both panels reveal some non-linearities located at different femaleness rates in the two countries. The Canadian dip is located around the 55 percent rate, while the American dip is located around the 80 percent rate. These differences are reflected in the estimates of g by type of job. In the United States, the correlation between log wages and *PFEM* changes monotonically as we move across jobs. For females, that largest penalty to *PFEM* is in male jobs, while the smallest is in female jobs. The opposite pattern is observed for males. Here the largest penalty is in female jobs while the smallest is in male jobs. Differences in the relative position of occupations will become an important ingredient in our account of Canada/U.S. differences in the correlation of wages with gender composition.

[25] In Canada, the highest provincial minimum wage (Ontario and Quebec's) was CA$4.75 (U.S.$5.78). In the U.S. the federal minimum wage was U.S.$3.35, but 10 states had higher minimums which ranged from $3.55 to $4.33.

[26] 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 for Canada and 0.065 for the United States.

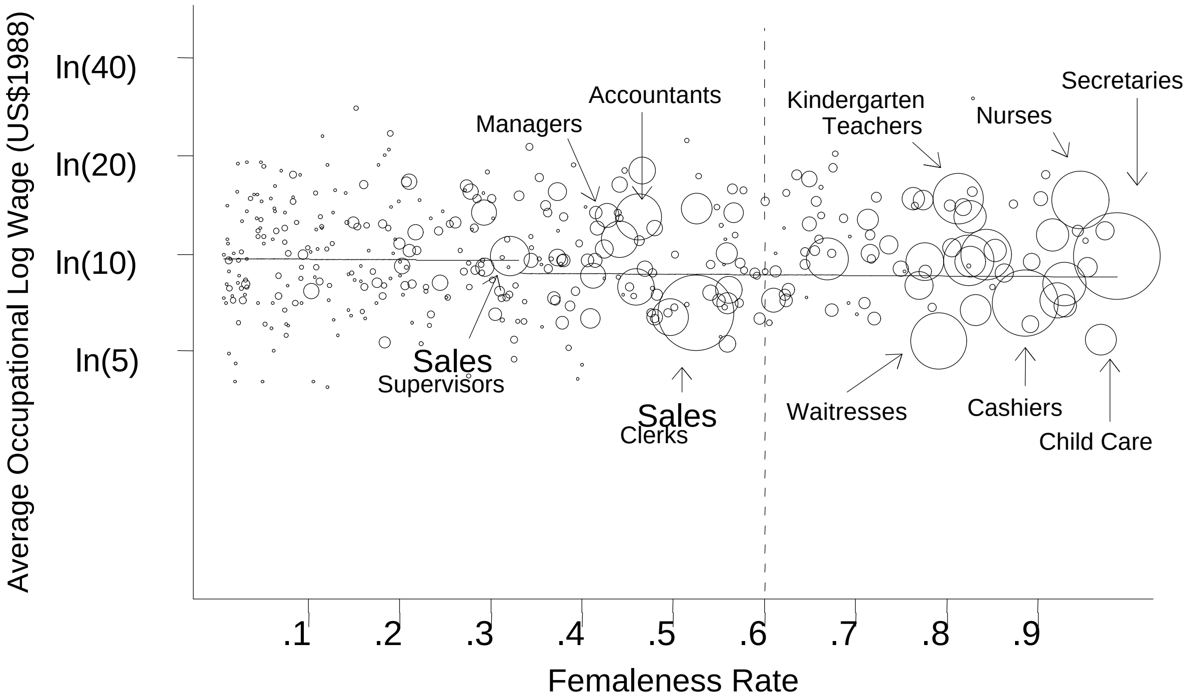

Figure 1. Effect of Gender Composition on Average Female Wages in Canada – 1988

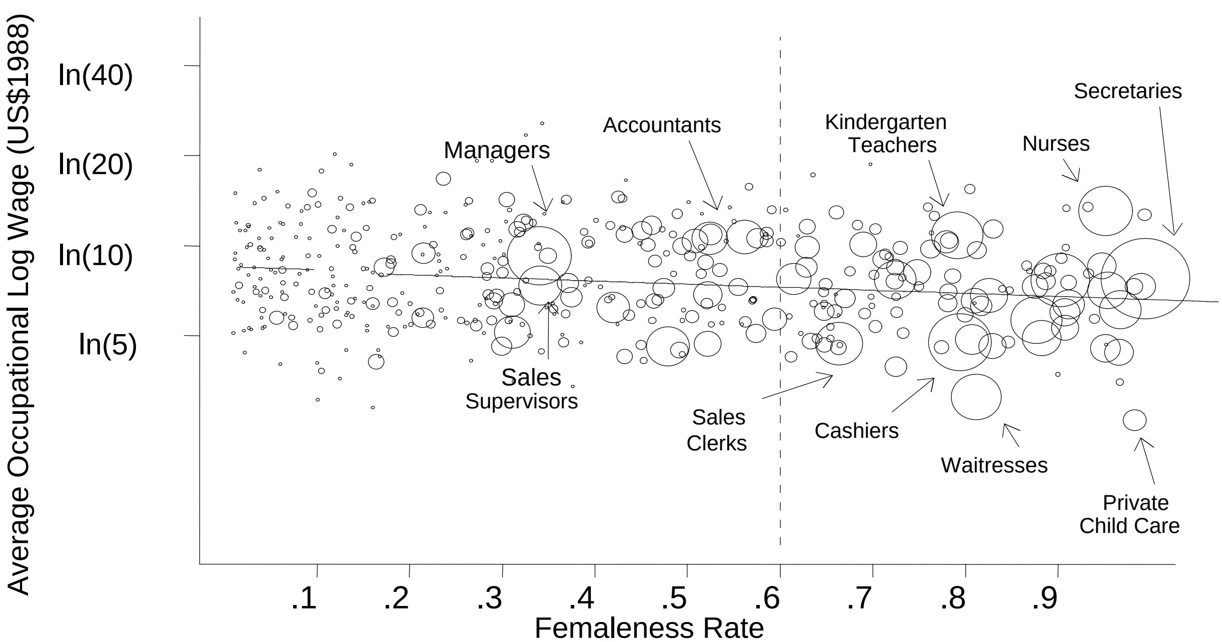

Figure 2. Effect of Gender Composition on Average Female Wages in the United States – 1988

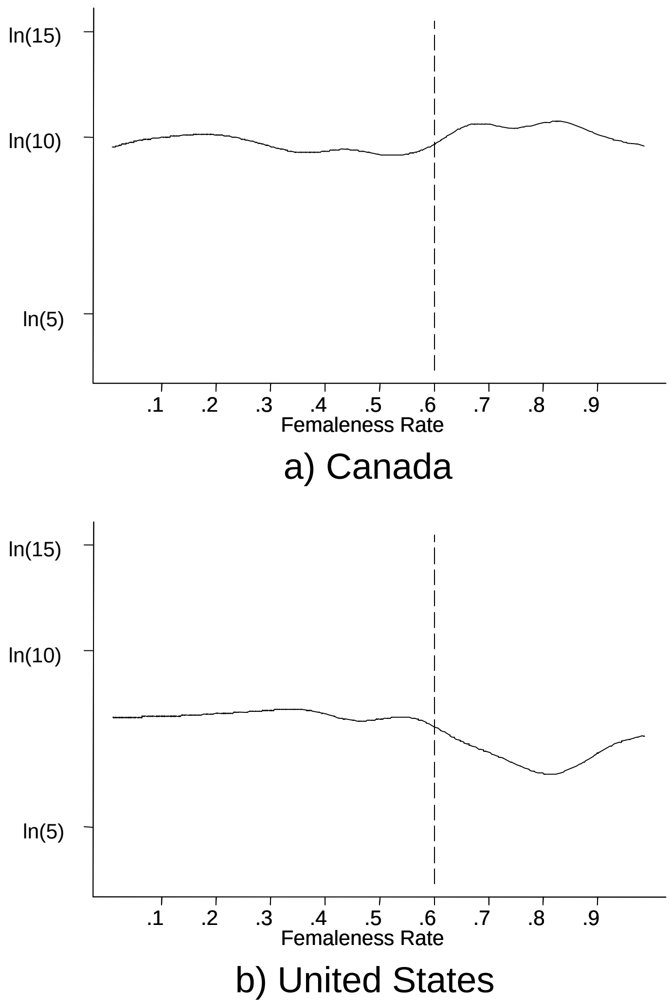

Figure 3. Weighted Kernel Regression of Average Occupational Female Wages on Gender Composition – 1988

Finally, in table 1 we also report the unadjusted female/male wage ratio, which averages 76 percent in Canada (for all jobs) and 72 percent in the United States. It is consistently higher in Canada, although the cross country difference is not substantial.[27] 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

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

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 ratio by about 1 percentage point in both countries.

The education variables in the LMAS do not record years of education, which is available in the CPS-ORG. Using the U.S. years of education and the "final year completed" variables, we were able to classify the U.S. data into six education classes largely comparable to those available in the LMAS. The percentages of women and men in each educational category, along with the means of other variables for the Canadian and U.S. samples in 1988 are reported in table 2. The U.S. samples show higher average levels of education, seen most clearly in the percentages with only a primary education and with a university degree.

Americans are also more likely to be non-white, reinterpreted here as members of a visible minority. The coding of the "visible minority" variable in Canada is, however, a subject of controversy. It is a constructed variable from data on ethnic background and is likely to also capture immigrant status, and therefore cannot be readily compared with the American variable. As a consequence, we do not emphasize Canadian-American differences in this dimension.[28]

There is generally less than one percentage point of difference in the distribution of workers by industrial sectors between the two countries. The exceptions are durable manufacturing and trade which groups 1.5 percent and 3 percent more workers, respectively, in the United States than in Canada, and public administration which groups 2.5 percent more workers in Canada than in the United States. This last difference is not as high as might be expected. One should also note that in both countries, about 30 percent of women work in the "public goods" sector: medical, welfare and educational services. Differences between the two countries in consumer

[28] We have investigated the contribution of race to the relationship between wages and femaleness rates in the United States. We estimated our regressions using a sub-sample of white Americans and found no substantial differences from the results using the complete sample. For example, the "raw" regression estimates are -0.234 for females and 0.001 for males.

Table 2
Means of Selected Variables – 1998

	Women		Men	
Variable	Canada	U.S.	Canada	U.S.
Wage (1988 U.S.$)	8.95	8.35	11.69	11.51
St. Dev. of Wages	(4.56)	(5.64)	(5.60)	(6.91)
Age	36.5	37.2	37.2	37.3
Education:				
Primary	.063	0.33	.104	.056
Some High School	.101	0.87	.130	.111
High School Graduate	.362	.404	.341	.362
Some Post-Secondary	.101	.115	.097	.096
Post-Secondary Degree	.210	.141	.162	.126
University Degree	.164	.220	.167	.248
Part-time	.226	.168	.042	.046
Married	.665	.569	.690	.646
Visible Minority	.052	.152	.051	.132
Metropolitain Area	.731	.802	.703	.800
Industrial Sector:				
Agriculture, Forestry and Fisheries	.011	.007	.023	.022
Mining	.006	.003	.029	.011
Construction	.017	.013	.085	.099
Manufacturing				
Nondurable	0.73	.077	.110	.093
Durable	0.47	.074	.159	.175
Transportation and public utilities	0.46	0.45	.116	.106
Trade	.161	.195	.156	.178
FIRE	.088	.096	.040	.049
Business and professional services	.062	.079	.043	.081
Consumer services	.121	.060	.055	.028
Medical, welfare, and educational services	.291	.301	.098	.098
Public administration	.075	.051	.086	0.60
Federal	0.20	.016	.042	.019
Provincial (State)	0.29	.018	.023	.016
Local	.016	.016	.035	.025
Union coverage	.371	.157	.452	.236
Tenure	5.78		8.00	
Establishment Size:				
s < 20	.376		.300	
20 <= *s* < 100	.298		.320	
100 <= *s* < 500	.203		.237	
s >= 500	.122		.142	
Nº. of observations	14,868	76,979	17,739	84,009

services and business services should be de-emphasized as the classification of basic industries into these aggregates can differ across countries.[29] Similarly, the Canadian federal sector includes the main industries that are under federal jurisdiction and is not directly comparable to the corresponding U.S. sector.

One dramatic difference between the two countries is the proportion of workers covered by collective bargaining. Union coverage rates in Canada are almost double the U.S. rates. These differences in the unionization rates have been studied in detail elsewhere. 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 coverages rates of 45.8 percent for males and 36.4 percent for females. Our rates are marginally higher than Riddell's and roughly similar to Lemieux's (45.2 percent for males and 37.1 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. Note that our rates, as well as those of Lemieux and Riddell, are higher than those reported by Doiron and Riddell (1994) for 1988 LMAS (38 percent for males and 29 percent for females).

An illustration of the potential impact of unionization on the effect of gender composition on female wages is shown in figure 4. Figure 4 plots the kernel density estimates, which can be understood as smoothed histograms, of female wages by job types in the two countries.[30] The union coverage rates among women in 1988 are 43 percent for female jobs, 26 percent for mixed jobs, and 35 percent for male jobs in Canada. In contrast, union coverage among women

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[29] For example, photographers and travel services are classified as consumer services in Canada. In the United States, those industries do not appear in the 3-digit industry codes. It is thus not possible to know where they are classified.

[30] Kernel density estimates are easily understood by reference to histograms. Histograms represent the frequencies of observations in a number of bins of a given width, which determines the smoothness of the histogram With kernel density estimation, a similar parameter is called bandwidth; here a bandwidth of 0.07 is used. In an histogram, the frequency of observations in any given bin can be computed as the number of observations times a rectangular weighing function of given bin-width. Instead of using a rectangular weight function, the kernel density estimates presented here use a Gaussian weight function, called the kernel, and can be characterized as a sum of `bumps' placed at the observations. Note that 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 types of jobs.

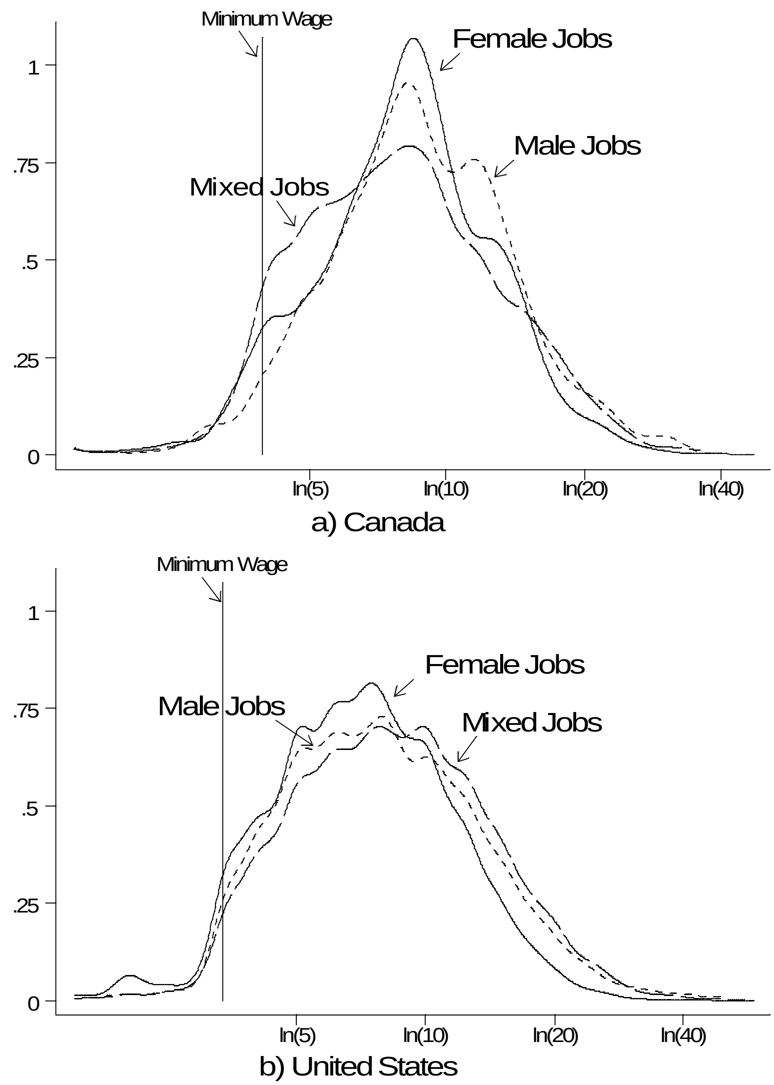

Figure 4. Density Estimates of Female Wages - 1988

decreases with the femaleness of employment in the United States, the corresponding rates for the female, mixed and males jobs are 15 percent, 16 percent and 19 percent.[31] As argued in DiNardo, Fortin and Lemieux (1996), unionization leads to a more compressed wage structure. Correspondingly, the densities of female wages in both female jobs and male jobs in Canada share *the same mode* and are much more compressed than the corresponding densities in the United States. Doiron and Riddell (1994) argue that the gender wage gap would have increased 7 percentage points between 1981 and 1988 if not for the reduction in the gender unionization

[31] Further comparisons of cross-country differences in unionization rates by jobs are done in section 6.

gap which occurred over this period. We will thus examine the potential contribution of differences in unionization rates to cross country differences in the correlation of wages and *PFEM* below.

Finally, our Canadian samples have a few additional variables, such as tenure and establishment size, which we use in some parts of the analysis. Males in Canada have greater tenure than females and are more likely to work at large establishments.

Differences in the occupational characteristics of the jobs in which women and men work have 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 provide a complete view of the Canadian evidence, 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.[32] The job characteristics are available for the seven-digit occupations codes (more than 6,500 categories) and, in the absence of appropriate weights, have to be averaged over the four-digit categories.[33] 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. Treiman (1979), Miller, Treiman, Cain and Ross (1980)).[34]

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[32] While Hunter and Manley 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 conditions.

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

[34] Treiman (1979) provides a discussion of the gender biases that may arise in job evaluation systems. He argues that the way many systems measure job characteristics (most commonly skill, effort, responsibility and environment) may favour male jobs, or permit greater differentiation among male jobs. For example, effort is often measured by strength requirements rather than levels of fatigue; manual skill focuses on the ability to work with tools rather than manual dexterity; responsibility is measured in terms of supervision rather than organization. He speculates that this gender bias may result from the industrial origins of many job evaluation schemes.

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

(1) $$\ln w_i = X_i b + a_k \cdot OCC_{ki} + n_i ,$$

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 n_i is an individual specific error term. The correlation of the occupation fixed effects, a_k, with the gender composition of that occupation, which is our primary interest, is specified as

(2) $$a_k = l + g\, PFEM_k + h_k ,$$

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

(3) $$\ln w_i = l + X_i b + g\, PFEM_k + \left(h_k + n_i\right).$$

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 h_k.[35]

One way to proceed would be to estimate (3) directly by generalized least-squares (GLS). An alternative is the following two-step procedure.[36] First, estimate equation (1) by OLS, or in our case weighted least-squares (WLS) as we use the LMAS or CPS supplied individual level weights in the estimation. We can express the resulting estimates of the occupation effects as

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

[36] Amemiya (1978) compares the properties of the one-step GLS estimator and two-step estimators. If we use GLS in the second stage there is exact equivalence. Of course, in application we use feasible GLS in the second stage which is based on a particular assumed structure of the error term.

(4) $$\hat{a}_k = a_k + ,_k ,$$

where $,_k$ is the measurement error in the $\hat{a}_k$. We then estimate the equation

(5) $$\hat{a}_k = l + g\,PFEM_k + (,_k + h_k),$$

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 g. The appropriate estimation strategy for (5) depends on which error component, $,_k$ or h_k, dominates the composite error term. On the one hand, $,_k$ 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 a_k. On the other hand, there is no obvious reason why h_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.[37]

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 $\hat{a}_k$ from the first stage regressions as weights.[38] In GLS2 the sum of the LMAS or CPS 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 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 the OLS estimates.[39]

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[37] This strategy thus takes jobs as unit of observation rather than individuals. For problems with 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. Both the U.S. and Canadian 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 (262) male occupations, 133 (120) mixed occupations and 80 (115) female occupations in our Canadian (American) sample.

[38] Since the first stage regressions are estimated by weighted least-squares using the LMAS and CPS sample weights, following Wooldridge (1998) it might be preferable to use White estimates of the sampling variances of the $\hat{a}_k$ 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).

[39] 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 -.0986 with OLS to -.1305 with WLS.

5. Results

5.1 Adjusted 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 and the United States, progressively adjusting for individual level productivity characteristics in the first stage regressions. In the first row for each year we control for "human capital" variables: a quartic in age and six education classes.[40] The results confirm previous findings that the largest changes in the effect of the femaleness rate on wages with the inclusion of human capital variables are for males. In the second row for each year we add explanatory variables in an attempt to replicate the conditions in which a comparable worth 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 comparable worth 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 (Canada) or region (U.S.) effects, 11 industry effects and dummy variables for metropolitan area, employment in the federal, provincial/state or local governments, union coverage and part time status. The effects of this change in specification are smaller parameter estimates for each group. The larger changes are observed for American females and Canadian males.

In the last specification we add individual characteristics, some of which are unlikely to be considered legitimate bases of wage variation in legislation. These include tenure, establishment size, the numbers of preschool and older children respectively (up to 3) (for 1988) and dummy variables for marital status and visible minority status. Note that some of these variables are not available in the CPS and therefore only the estimates for Canada are presented. In each year, and

[40] The returns to these human capital variables are reported in table A-2 for 1998. They show the higher returns to education for U.S. males, found elsewhere in the literature.

for either gender, the effect of these new variables is very small. The estimates of g remain essentially unchanged.

Table 3
Canada–U.S. Comparison of the Effect of Occupational Femaleness on Wage Levels

Year	Canada			United States		
Specification:	UWLS	GLS1	GLS2	UWLS	GLS1	GLS2
1987: Women						
1: Human capital	-.146	-.091	-.004	-.307	-.273	-.212
	(.057)	(.052)	(.047)	(.052)	(.048)	(.050)
2: 1 + Sectoral	-.108	-.056	-.040	-.164	-.150	-.155
Controls	(.051)	(.045)	(.036)	(.048)	(.043)	(.043)
3: 2 + Individual	-.120	-.066	-.041			
characteristics	(.049)	(.043)	(.034)			
No. of occupations		380			449	
1988: Women						
1: Human capital	-.013	-.013	-.023	-.230	-.223	-.213
	(.060)	(.055)	(.046)	(.055)	(.048)	(.050)
2: 1 + Sectoral	-.037	-.012	-.066	-.101	-.124	-.164
Controls	(.054)	(.050)	(.037)	(.051)	(.044)	(.043)
3: 2 + Individual	-.033	-.012	-.062			
characteristics	(.051)	(.047)	(.035)			
No. of occupations		378			451	
1987: Men						
1: Human capital	-.207	-.229	-.217	-.269	-.284	-.148
	(.042)	(.040)	(.036)	(.043)	(.039)	(.048)
2: 1 + Sectoral	-.081	-.099	-.052	-.156	-.171	-.044
Controls	(.039)	(.031)	(.033)	(.041)	(.038)	(.045)
3: 2 + Individual	-.076	-.095	-.067			
characteristics	(.037)	(.034)	(.030)			
No. of occupations		473			493	
1988: Men						
1: Human capital	-.274	-.252	-.228	-.275	-.273	-.149
	(.042)	(.040)	(.038)	(.043)	(.041)	(.049)
2: 1 + Sectoral	-.159	-.141	-.100	-.155	-.154	-.042
Controls	(.039)	(.037)	(.034)	(.041)	(.039)	(.046)
3: 2 + Individual	-.151	-.131	-.110			
characteristics	(.037)	(.035)	(.031)			
No. of occupations		456			493	

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 or CPS) weights (by occupation) are used as weights. All the underlying first stage regressions are estimated by weighted least-squares using LMAS or CPS sample weights. Human capital conditions on a quartic in age and on six education classes. Sectoral controls add dummies for province (10) or region (9), metropolitan area, industry(12), employment in the federal, provincial or state, 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).

In attempting to summarize the results in table 3 it is necessary to reconcile any differences in the results across years, and in some instances across the different estimation strategies. We first discuss the results for men, which are in line with the rest of the literature, and then turn to the more controversial results for women.

First, controlling for age and education has substantial effects on our estimate of g for American men (second panel of table 3). Recall from table 1 that the "unadjusted" estimate of g for this group was roughly 0. In the Human Capital specification the average UWLS estimate is about -0.27, implying an elasticity of -0.068 at an average *PFEM* of 0.25. As noted by Macpherson and Hirsch (1995), the small estimate from the specification with no additional control variables is due to low skill, low pay, predominately male occupations. Once some control for skills is made, the estimate is much larger.

Note also that the results from the richer specifications for this group are generally consistent across years but not across the UWLS and GLS estimation strategies. The original discussion of these 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 are estimating 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 WLS estimates of the sampling variances of the $\hat{a}_k$ from the first stage regressions should be proportional to occupational sample size. In application, the GLS1 results are actually in greater agreement with the UWLS than the GLS2 estimates.

If g 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. In the present context, there is some evidence that g varies with occupation size. In table A-1 of the appendix we decompose the results for 1987 by decile of the sum of the individual weights (i.e., the weights used for GLS). For each decile we present a UWLS estimate of g. The estimates are uniformly negative except the result for the largest occupations which is positive (although statistically insignificant). This is the estimate, however, which receives the largest weight in the GLS2 estimation. Therefore the GLS2 results for American males can be viewed as reflecting the fact that conditional on individual characteristics, the average male faces a modest penalty due to the virtual absence of a penalty in large occupations.

The major discrepancy in the results for Canadian males is in the estimates across years. In the richer specifications, the 1987 results are generally one half their 1988 counterparts using the UWLS estimation strategy. A limitation of the Canadian data is that the smaller sample sizes mean that the same occupations are not necessarily observable in both years, and for those that are that the estimate of mean wages can change dramatically. The first problem is clearly evident for Canadian males 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 $\hat{g}$ in 1987 and -0.150 (0.037) in 1988. A second consideration is that the 1987 results are sensitive to a few observations.[41] Simply excluding four influential but small occupations leads to an estimate of g of -0.114 (0.036) 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 $\hat{g}$ = -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 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 g for Canadian males.

[41] A useful measure of the influence of an observation is the DFBETA which measure the difference between the regression coefficient, here $\hat{g}$, when the *i*th 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 $2/\sqrt{n}$.

We next consider the results for women. For American females, reconciling the results from the different specifications across estimation strategies is an easy task. Using the second specification as a basis of comparison, there is consistent evidence that $\hat{g}$ is about -0.14 for these women.

Perhaps the most important and potentially controversial reconciliation is for Canadian 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 3 as a basis of comparison, the UWLS estimate of g for 1987 using the common occupations is -0.083 (0.048) and for 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.[42] Excluding this occupation from the 1988 sample leads to an UWLS estimate (specification 3) of -0.055 (0.050). 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.

These conclusions in turn point to some interesting Canada/U.S. differences in the penalty for women, although there is some sensitivity to how the comparison is made. On one hand, the simple differences between the point estimates for the two groups are at best marginally significant.[43] On the other hand, there is little consistent evidence that Canadian females face a penalty to working in female jobs.

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, for example American males, the straightforward

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[42] This conclusion was reached examining the DFBETAs.

[43] Given the estimates come from independent samples, the standard error of the difference is just $\sqrt{Var(g_{CA}) + Var(g_{US})}$.

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 correlation for females and one-half the correlation for males is due to these sorts of factors. Furthermore, they argue that once control for detailed occupational characteristics is made, the correlation 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 4) we start from the final row of table 3 and add controls for the CCDO skill requirements characteristics: general educational development (GED), measured in approximate of years of schooling, and specific vocational preparation (SVP), measured in months of training. In Canada, controlling for skill requirements decreases the magnitude of g 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.[44] Finally, in specification 7 we add controls for outside and inside work, corresponding to the CCDO work location variable (EC-1).

[44] 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.

Table 4

The Role of CCDO Occupational Characteristics in the Effect of Gender Composition on Wages in Canada – 1998

		Women	Men
4:	3 + Educational requirements[a]	-0.11	-.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.

Overall, these additional controls lead to an estimate of g 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.

5.3 Gender Composition Coefficients among Alternative Worker Groups

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 g we obtain for Canadian females result from these sorts of differences; if we focus on full time workers we may recover the "expected" larger negative estimates. Finally, in Canada the wage structure is known to favour older workers (Morissette and Bérubé (1996), Beaudry and Green (1997)). In the United States the wage structure works to the advantage of more educated workers, as shown by the increase in returns to education over the last 15 years (Katz and Murphy (1992)), an increase not witnessed in Canada (Bar-Or, Burbidge, Maggie and Robb (1995)).[45] Therefore, decomposing the results by age or education may also be of interest. We restrict our analysis to females, as this is the group that is typically the target of pay equity legislation.

In table 5 we present estimates of g for females in Canada and the United States (in 1988) by these different groupings.[46] The results tend to support our aggregate inference, but there are some interesting exceptions. In both countries $\hat{g}$ tends to be larger in nonunion and full time employment, and among university graduates. It is difficult to compare the estimates among nonunion workers in Canada and in the United States. Some occupations in particular teaching occupations and health care occupations are almost completely unionized in Canada, and they will be virtually excluded in a regression using nonunion workers in Canada but not in the corresponding regression for the United States. We prefer to compare coefficients estimated across the same occupations in the two countries. We provide a framework for doing so in Section 6.

It has also being suggested that the negative effect of gender composition on Canadian female wages may be larger among particular groups of workers. When we restrict our attention to the sub-sample of full-time non-unionized women (47 percent of working women), we find estimates of g ranging from -.236 to -250 (with standard errors 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 g ranging from -.315 to -.336 (with standard errors around 0.1).

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[45] In our cross-sectional analysis, the latter cross-country difference in returns to education for men is illustrated in table A-2.

[46] Unfortunately, there is no Canadian variable equivalent to the "class" variable of the CPS that distinguishes workers by public/private sector status. The variable used in Riddell (1993) for 1986 jobs has not been coded for any other labour force survey. Note however that the estimates for females working in the public/private sectors are very similar in the United States: -.229 (.071) in the public sector vs -.249 (.062) in the private sector.

Table 5

Gender Composition Coefficients on Female Wages Among Alternative Worker Groups – 1998

Specification: Group	*NC*	No controls	(1) Human Capital	(2) 1+Sectoral Controls	*NC*	No controls	(1) Human Capital	(2) 1 + Sectoral Controls
			Canada:				*United States:*	
Age:								
16-29	307	-.075	-.045	-.057	395	-.256	-.200	-.171
		(.061)	(.049)	(.041)		(.063)	(.052)	(.044)
30-44	307	-.059	-.050	-.109	410	-.241	-.241	-.174
		(.071)	(.058)	(.047)		(.064)	(.052)	(.046)
44-69	246	.102	.073	.009	384	-.154	-.169	-.117
		(.079)	(.064)	(.055)		(.067)	(.057)	(.050)
Education:								
Drop-out	230	-.113	-.114	-.087	308	-.318	-.299	-.197
		(.060)	(.059)	(.052)		(.048)	(.046)	(.039)
High School	294	-.028	-.018	-.032	389	-.158	-.149	-.107
		(.052)	(.048)	(.038)		(.051)	(.049)	(0.40)
Post-	260	.045	.045	-.001	354	-.202	-.190	-.145
Secondary		(.063)	(.058)	(.049)		(.062)	(.056)	(.049)
University	179	-.095	-.120	-.184	328	-.315	-.350	-.272
		(.081)	(.075)	(.066)		(.057)	(.055)	(.052)
Union coverage status:								
Nonunion	342	-.182	-.142	-.136	439	-.254	-.224	-.186
		(.059)	(.048)	(.042)		(.063)	(.052)	(.045)
Union	287	.025	.010	.044	358	-.004	-.094	-.038
		(.061)	(.047)	(.060)		(.058)	(.044)	(.042)
Hours status:								
Part-time	211	.353	.323	.169	303	.033	.012	-.010
		(.099)	(.083)	(.066)		(.092)	(.079)	(.071)
Full-time	373	-.097	-.082	-.107	449	-.227	-.209	-.176
		(.058)	(.043)	(.035		(.058)	(.046)	(.041)

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 (i.e., GLS2). Estimated standard errors are in parentheses. *NC* is the number of occupations

In summary, while there are particular groups of Canadian women who face a negative penalty for working in female occupations, our general conclusions continue to hold. Not all pair-wise comparisons result in statistically significant differences, however the overall pattern of coefficient estimates suggest a stronger negative effect of the femaleness of occupations on female wages in the United States than in Canada.

6. Accounting for Canada-U.S. Differences in the Effect of Gender Composition on Female Wages

To determine if the Canada-U.S. differences in g we observe are an artifact of sample sizes, differences in variable coding, etc., or, rather the result of actual differences in wage structures we provide a direct investigation into their sources. A first step to this goal is to use the same occupation codes in the two countries. As explained in Section 3, we construct an occupational crosswalk between the Canadian and U.S. codes, which reduces the number of possible occupation categories to a maximum of 310. In the first two rows of table 6, we report estimates of g for females in Canada and the United States using these new codes. In most cases, the estimates are marginally smaller than their counterparts in table 3.[47]

Table 6
Accounting for Canada–U.S. Differences in the Effect of Gender Composition on Female Wages – 1998

Simulation	Specification:	No controls	(1) Human capital	(2) 1 + Sectoral Controls
0:	Canada using occupational cross-walk	-.022	-.019	-.060
		(.070)	(.053)	(.042)
1:	United States using occupational cross-walk	-.192	-.179	-.136
		(.077)	(.061)	(.051)
2:	1 + Canadian variance	-.176	-.164	-.124
		(.070)	(0.56)	(.047)
3:	1 + Canadian unionization structure	-.156	-.158	-.131
		(.078)	(.061)	(.051)
4:	2 + Canadian unionization structure	-.143	-.145	-.120
		(.072)	(.056)	(.047)
5:	1 + Canadian ranking of occupations	-.075	-.061	-.019
		(.079)	(.062)	(.055)
6:	3 + Canadian ranking of occupations	-.034	-.035	-.009
		(.082)	(.064)	(.055)

Note: Estimated standard errors are in parentheses. They do not take into account errors from the simulation experiments and should be viewed as lower bounds.

[47] In a related experiment we substituted Canadian femaleness rates for the American ones. This led to larger (in absolute value) estimates of g .

An often discussed difference between the Canadian and U.S. wage structure is in the returns to skills, which increased substantially in the United States during the 1980's. In table A-2 we report the estimated parameters on the explanatory variables in our specification 1 (estimated with the original occupation codes). We see large Canada-U.S. differences in the returns to education for males but not for females. For women, returns to human capital are virtually identical in the two countries, once we control for occupations. To assess the role of cross-country differences in the returns to skill, we examine the correlation between female wages and the femaleness rate in the United States when women there face the Canadian returns to human capital. More precisely, we apply our estimation strategy to log wages predicted by

$$\ln w_i^{US} = \hat{b}^{CAN} X_i^{US} + \hat{a}_k^{US} \cdot OCC(k)_i^{US} + \hat{n}_i^{US}. \tag{6}$$

Not surprisingly, we do not find any difference in our estimate of g (and do not report it), and conclude that differences in returns to observable skills, (or rather the absence of differences) can not account for cross country differences in the effect of gender composition.

Following Juhn, Murphy and Pierce (1993), increases in the returns to unobserved skills have been offered as a source of cross-country differences in the gender wage gap (Blau and Kahn 1998). To conduct a simulation that asks what the correlation between female wages and the femaleness rate would be in the United States if the dispersion of returns to unobserved skills were more compressed as in Canada, we would have to use the following predicted log wages

$$\ln w_i^{US} = \widehat{b^{US}} X_i^{US} + \hat{a}_k^{US} \cdot OCC(k)_i^{US} + \hat{n}_i^{US} \cdot \left(\hat{s}_n^{CAN} / \hat{s}_n^{US}\right), \tag{7}$$

where $\hat{s}_n^{C}$ is the standard deviation of the residuals from the corresponding regression in the indicated country. However since this does not affect the estimated occupation fixed effects used in the second step, this simulation is ineffective in our econometric framework. Rather, we simply normalize the distribution of U.S. log wages so that their estimated standard deviation is equal to its (estimated) Canadian counterpart. The resulting estimates of g are reported in row 2 of table 6. They suggest that decreasing the U.S. standard deviation of log wages accounts for at most 10 percent of the Canada-U.S. difference in the coefficient on *PFEM*. Overall, these simulations suggest that explanations of cross-country differences in the relative economic stature of the genders based on corresponding differences in the returns to observed and unobserved skills have little explanatory power for the Canada-U.S. differences here.

A striking Canada-U.S. difference, mentioned in Section 3, is in union coverage rates. The differences in unionization rate by job types among women, noted earlier (with 43 percent of women in female jobs being unionized in Canada vs. 15 percent in the United States) become even more important comparing finer groups of occupations from our occupational crosswalk. Two important female occupations figure predominantly in this comparison: health care workers (approximately 10 percent of female workers) and teachers (approximately 5 percent of female workers). In Canada health care workers have very high rates of unionization (e.g., more then 85 percent among nursing and therapy occupations, around 60 percent among technologists), while in the United States unionization rates in those occupations is less then 20 percent. Among elementary and secondary teachers, union coverage for women is close to 90 percent in Canada while it is only 60 percent in the United States; among post-secondary teachers, the percentages are 75 percent vs. 25 percent. Large differences in unionization rates are also observed for less important occupations. For example, the Canada-U.S. differences are: 50 percentage points for Food and Beverage Preparation Occupations n.e.c. (1 percent of female workers), 46 percentage points for Personnel and Related Officers (0.5 percent of female workers), 39 percentage points for Librarians, Archivists and Conservators (0.5 percent of female workers).

To simulate the Canadian union coverage in the United States, we take advantage of the fact that our data carry sample weights and use a reweighting procedure in the spirit of DiNardo, Fortin and Lemieux (1996). Let f_i^{US} denote the U.S. sample weight of observation *i* and let *u* be a dummy variable that takes on the value 1 if individual *i* is covered by collective bargaining and the value 0 if not. To simulate the Canadian unionization structure, we replace this weight by

$$f_i^*(u) = \begin{cases} f_i^{US} \cdot (y_{u|x}^{CAN}(u,x) / y_{u|x}^{US}(u,x)) & \text{if} \quad u = 1, \\ f_i^{US} \cdot ((1 - y_{u|x}^{CAN}(u,x)) / (1 - y_{u|x}^{US}(u,x))) & \text{if} \quad u = 0, \end{cases} \tag{8}$$

where $y_{u|x}^{C}(u,x)$ is the reweighting function of country *C*. An estimate of the reweighting function $y_{u|x}^{C}(u,x)$ can be obtained by estimating the conditional probability $\Pr(u = 1 \mid x, C)$ using the probit model

(9) $$\Pr(u = 1 \mid x, C) = \Pr(\epsilon > -b^{C} H(x)) = 1 - N(-b^{C} H(x)),$$

where *N*(.) is the cumulative Normal distribution and *H(x)* is a vector of covariates that is a function of *x*. We specify the vector *H(x)* as a quartic in age, six education classes, 11 industry effects, and dummy variables for federal, provincial (state) or local government employment, metropolitan area, marital status, and part-time status. Row 3 of table 6 shows that differences in union coverage account for a modest proportion of the Canada-U.S. difference, and are ineffective when industry controls are introduced (specification 2). Combining differences in union coverage with differences in the dispersion of log wages can account for up to a 20 percent of the cross country gap (row 4), but again there power is reduced in specification 2.[48]

Another salient difference between the two countries is the relative position of the different job types. These differences are clearly illustrated in figure 5, which superimposes the kernel density estimates of the distribution of the log wages of women and men by job types. Particularly striking is the panel that displays the density of female wages in female jobs. The U.S. density is everywhere to the left of the Canadian density. The Canadian distribution has greater mass between $5.00 and $8.00 suggesting that more than a higher minimum wage is at play.[49] For mixed jobs, the reverse is true. To simulate the Canadian ranking of occupations in the U.S. wage structure, we begin by ranking the occupations in the overall distribution of wages (women and men combined). That is, each wage level is assigned a rank in the overall wage distribution and the rank of an occupation is computed as the average rank of each woman or man in that occupation. These average ranks for women and men, along with the median ranks, are reported in table 7. There we see that while average ranks for women and men on all jobs are about the same in the two countries, their distribution across job types is very different. In particular, workers in mixed jobs in the United States are positioned at a higher percentile than workers in other jobs. This pattern is also apparent from the middle panels of figure 5.

[48] Increasing the union coverage rates in the United States may not fully capture the impact of unionization. As union density declined dramatically in the Unites States over the 1980's, unions also lost some of their ability to compress wages. When an alternative experiment is conducted for Canada; that is, lowering union coverage rates to the American ones, the raw correlation rises to -0.0989, explaining 36 percent of the cross-country difference.

[49] Alternatively, important spill-overs of the minimum wage could be at work. However, we do not investigate this issue. Note that a similar pattern is seen for female wages in male jobs. However, these account for less than 10 percent of female workers.

Table 7
Canada – U.S. Comparison on the Ranking of Occupations in the Overall Wage Distribution and Within Occupation Wage Gap by Job Types

	Women			Men			
Sample	No. of Occupations	Average Centile	Median Centile	No. of Occupations	Average Centile	Median Centile	Within Occupation Wage Gap
Canada: 1988							
All jobs	277	40.6	39.4	310	57.4	60.3	.226
Female jobs	65	41.2	40.8	63	56.2	56.1	.143
Mixed jobs	83	39.1	35.5	83	56.5	59.0	.248
Male jobs	129	42.5	39.1	164	58.1	62.8	.283
United States: 1988							
All jobs	293	41.3	44.3	309	57.1	59.2	.219
Female jobs	71	38.6	42.1	71	47.9	53.0	.179
Mixed jobs	81	46.1	50.1	81	61.7	62.2	.280
Male jobs	141	44.5	45.0	157	55.8	59.9	.198

Note: The rankings of occupations are computed with respect to the distribution of wages of both women and men in the specified country. The occupation categories are obtained from a cross-walk between the detailed occupation codes of each country, thereby aggregating the original 500 or so categories into a maximum of 310.

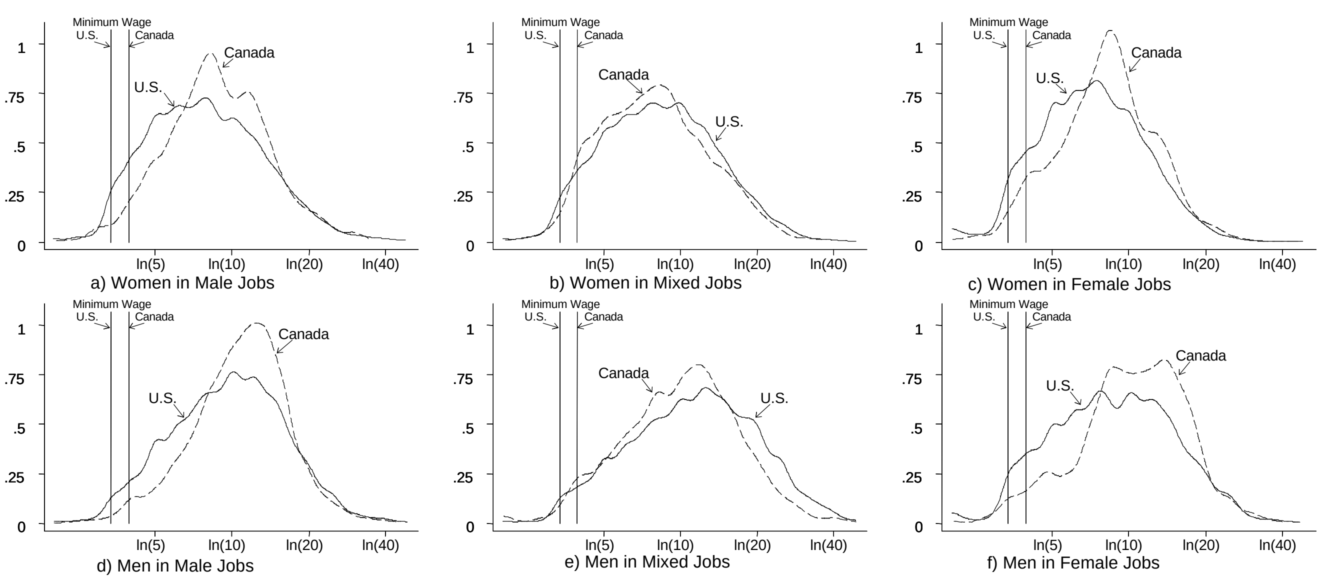

Figure 5. Density Estimates of Log Hourly Wages – 1988
(US$ 1988) hours weighted

Figure 6 shows the relative position of women's occupations in Canada and in the United States. We plot the Canadian rank of each occupation (using the occupational crosswalk) against the U.S. rank. Occupations that are on or around the 45 degree line, which is also drawn, rank similarly in the two countries. Occupations above this line, such as teaching occupations, nursing assistants, and social workers, rank higher in Canada. The relatively low ranking of teaching occupations in the United States is consistent with the industry-wage effects estimated by Helwege (1992). She finds that educational services industry-wage effects have steadily declined in the United States since the 1940s and were the second lowest in 1980.[50] Occupations below this line, such as managers, financial officers and sales managers, rank higher in the United States. This reflects the relatively higher position of mixed occupations in the United States.

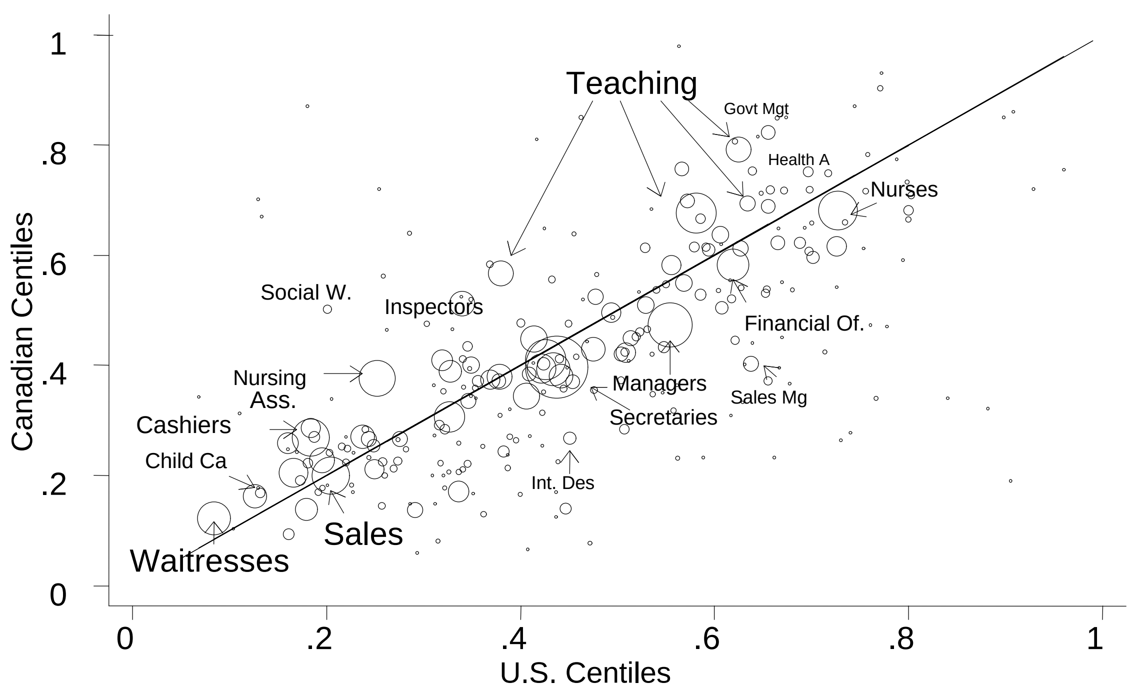

Figure 6. Ranking of Women's Occupational Wages in the Overall Wage Distribution – 1988

Let $p_{ki} = F^C(\ln w_{ki})$ be the position of woman *i* holding occupation *k* in the overall cumulative distribution of wages (women and men combined) $F^C(\ln w)$ of country *C*, and let

[50] Admittedly, these industry-wage effects are computed from a sample of white males!

$p_k^C = \sum_{i\in K} p_{ki} = \overline{F^C(\ln w_{ki})}$ be the average position of females in occupation *k* in country *C*. The occupational wage that an American woman in occupation *k* would have earned if her occupation had ranked as in Canada but if the U.S. wage structure prevailed is given by $\ln w_k^{CAN} = (F^{US})^{-1}\left[p_k^{CAN}\right] = (F^{US})^{-1}[\overline{F^{CAN}(\ln w_{ki})}]$. We simulate the wage of individual *i* by adding the difference resulting from the change in the positions of the average occupational wage $(\ln w_k^{CAN} - \ln w_k^{US})$ to her own wage

(10) $$\ln w_{ik5}^{US} = \ln w_{ki}^{US} + (\ln w_k^{CAN} - \ln w_k^{US}).$$

For example, secondary teachers, which are 47 percent female in Canada and 56 percent female in the United States, are ranked at the 80th percentile of the overall wage distribution in Canada and at the 62nd percentile in the United States. Since the U.S. log wages corresponding to the 62th and 80th percentile are 2.31 and 2.62, respectively, to simulate the increase from the change in relative position, we add a premium of 0.31 to the individual log wages of secondary teachers. The impact of these changes in relative position on the U.S. correlation between female wages and the femaleness rate is dramatic (row 5). They account for roughly 67 percent of the Canada-U.S. difference in specification 1 and almost all of the difference in specification 2. Also, adding in the adjustment for differences in unionization rates (row 6) further reduces the estimate of g in specification 1.

We conclude that unionization and occupation-industry wage-effects are the more important factors accounting for the Canada-U.S. difference in the effect of gender composition on female wages. In particular, a low female unionization rate in the United States and low occupation-industry wage-effects for "public good" sectors such as educational services work to the detriment of U.S. women.[51]

[51] Helwege (1992) has identified negative industry-wage effects in the government sector and the medical services sector, as well.

7. 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

$$\text{(11)} \qquad \overline{\ln w^j} = \hat{b}^j \, \overline{X^j} + \hat{a}_k^j \cdot \overline{OCC_k^j},$$

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

$$\text{(12)} \qquad (\overline{\ln w^M} - \overline{\ln w^F}) = (\hat{b}^M \, \overline{X^M} - \hat{b}^F \, \overline{X^F}) + (\hat{a}_k^M \cdot \overline{OCC_k^M} - \hat{a}_k^F \cdot \overline{OCC_k^F}).$$

The second term on the right hand side of (12) 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

$$\text{(13)} \qquad \overline{\hat{a}^j} = \hat{l}^j + \hat{g}^j \cdot \overline{PFEM^j}.$$

A standard Oaxaca decomposition of the second stage equations yields

$$\text{(14)} \qquad (\overline{\hat{a}^M} - \overline{\hat{a}^F} = (\hat{l}^M - \hat{l}^F) + \hat{g}^M (\overline{PFEM^M} - \overline{PFEM^F}) + \overline{PFEM^F} (\hat{g}^M - \hat{g}^F).$$

Equations (12) and (14) are related by noting that $\hat{a}_k^j \cdot \overline{OCC_k^j}$ in (12) is implicitly the sum $\sum_{l=1}^{K} \hat{a}_l^j \cdot \overline{OCC_l^j}$, and that $\overline{\hat{a}^j} \sum_{l=1}^{K} \hat{a}_l^j \cdot \overline{OCC_l^j}$ when we use GLS2 to estimate the second stage regression. Therefore, under the GLS2 weighting scheme equation (14) 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 (13) that

(15) $(\hat{g}^M \cdot \overline{PFEM^M} - \hat{g}^F \cdot \overline{PFEM^F}),$

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

One way of viewing (15) 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 $g^M = g^F = 0$).[52] Estimates of (15) are easily constructed for 1988 using average *PFEM* from table 1 and the GLS2 estimates of g^j for this year from table 3. For the United States the estimates range from 0.10 to 0.14 for the three specifications of *X*.[53] Given a gender log wage gap of 0.31 in this year, we see that approximately *one-third* of the gap is accounted for by the differences in g and *PFEM* across the genders. For Canada, the estimates range from -0.04 to 0.02.[54] Here the aggregate effect of g and *PFEM* is to lower the wage gap. 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 g. Since the log wage gap in Canada was 0.27 in 1988, these results suggest that policies aimed at eliminating the effects of gender composition would have limited effect on the log wage differential.

Following previous studies, in table 8 we present the Oaxaca decomposition's represented by (14). 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 $\hat{g}^M$, and that the GLS2 estimates of this parameter (table 3) are smaller than both Macpherson and

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

[53] The estimates are 0.145, 0.103 and 0.099 for specifications 1, 2 and 3 respectively.

[54] The estimates are -0.0181, -0.0419 and 0.0187 for specifications 1, 2 and 3 respectively.

Hirsch's result and the GLS1 estimates.[55] 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.

Table 8
Comparison of Decompositions in the Gender Gap – 1988

	Specification	Canada	United States
Total log wage gap		.273	.307
0:	No Controls	.273	.307
	Total due to Occupation Effects	(.019)	(.022)
	Part due to $\Delta PFEM$	0.61	.011
		(.022)	(.028)
	Part due to Δl and Δg	.213	.296
		(.019)	(.036)
1:	Human Capital		
	Total due to Occupation Effects	-.416	-.047
		(.015)	(.017)
	Part du to $\Delta PFEM$	.095	.060
		(.016)	(.020)
	Part due to Δl and Δg	-.511	-.107
		(.021)	(.026)
2:	1+ Sectoral Contrtols		
	Total due to Occupation Effects	-.356	.311
		(.012)	(.015)
	Part due to $\Delta PFEM$	.044	.017
		(.014)	(.019)
	Part due to Δl and Δg	-.400	.294
		(.019)	(.024)

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.

[55] Note that Macpherson and Hirsch (1995) use a weighted average of the male and female estimates. As explained in Section 5, the difference is accounted for by the non-linearity of the *PFEM* effect across occupations distinguished by size.

8. Conclusion

Our cross country comparison of gender composition and wages has identified some intriguing Canada-U.S. similarities and differences. Canadian males face a penalty for working in female jobs that is comparable to that faced by their counterparts in the United States. The story for females is much different. The estimated penalty for Canadian females is generally small and not statistically significant, while the penalty for American females is relatively large.

We attempt to account for the cross country differences in the penalties for females, examining corresponding differences in the returns to observable and unobservable skills, unionization and the ranking of different occupations. We conclude that both unionization and the relatively high occupation wage effects for certain public good jobs, such as educational services, work to the advantage of Canadian females.

Appendix

Table A-1
A Decomposition of the Correlation of Log Wages and Percentage Female by Decile of Occupation Size: Males – 1987

Decile	United States	Canada
First	-.047 (.191)	.066 (.166)
Second	-.562 (.175)	-.428 (.143)
Third	-.185 (.091)	-.113 (.118)
Fourth	-.409 (.121)	-.277 (.097)
Fifth	-.260 (.146)	-.286 (.111)
Sixth	-.369 (.086)	-.214 (.091)
Seventh	-.207 (.102)	-.202 (.086)
Eighth	-.276 (.101)	-.240 (.086)
Ninth	-.264 (.103)	-.247 (.098)
Tenth	.012 (.169)	-.238 (.147)

Note: "White" standard errors are in parentheses. The reported coefficients are OLS estimates of equation (5) from the sample of occupations lying in the indicated decile of the sum of the (individual level) sampling weights. The underlying individual level regressions include controls for education and age (specification 1 from Table 3).

Appendix (continued)

Table A-2
Effects of Human Capital Variables on Log Wage – 1988

	Women		Men	
Variable	Canada	U.S.	Canada	U.S.
Age	.168 (.031)	.153 (.013)	.220 (.028)	.166 (.013)
$Age^2 \div 100$	-.504 (.120)	-.467 (.050)	-.673 (.109)	-.486 (.051)
$Age^3 \div 10000$	.679 (.202)	.657 (.084)	.985 (.183)	-.726 (.086)
$Age^4 \div 1000000$	-.353 (.122)	-.357 (.050)	-.566 (.111)	-.444 (.051)
Education (High School Grad omitted):				
Primary	-.126 (.015)	-.114 (.009)	-.134 (.011)	-.219 (.007)
Some High School	-.060 (.011)	-.073 (.006)	-.070 (.010)	-.096 (.005)
Some Post-Secondary	.040 (.011)	.041 (.005)	.060 (.011)	.027 (.005)
Post-Secondary Degree	.094 (.009)	.087 (.005)	.084 (.009)	.054 (.005)
University Degree	.266 (.011)	.213 (.005)	.159 (.011)	.200 (.005)
Occupation Dummies	Yes	Yes	Yes	Yes
PFEM	-.013 (.055)	-.223 (.048)	-.252 (.040)	-.273 (.041)
No. of observations	14,868	76,979	17,739	84,009

Note: Standard errors in parentheses.

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