The evolution of male-female earnings differentials in Canadian universities, 1970–2001

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Abstract. The evolution of academic salaries and the gender earnings gap in Canadian universities is analysed using a unique Statistics Canada data set containing detailed information on all full-time teachers at Canadian universities over the period 1970 through 2001. Male salaries declined across cohorts, while female salaries remained stable; consequently, male-female earnings differentials have narrowed. Recent years have also seen increasing variation in academic salaries for both men and women. A gender earnings gap remains, the bulk of which can be explained by differences in men's and women's rank and field. The unexplained earnings gap has decreased over time. JEL classification: J16, J31

Evolution des différentiels de salaires entre hommes et femmes dans les universités canadiennes 1970-2001. On analyse l'évolution des salaires et l'écart entre hommes et femmes dans les universités canadiennes entre 1970 et 2001 à l'aide d'une base de données unique de Statistiques Canada qui contient des renseignements détaillés sur tous les professeurs à temps plein des universités canadiennes. Les salaires des hommes ont décliné au fil des cohortes pendant que ceux des femmes demeuraient stables, ce qui fait que les différentiels de salaires se sont amenuisés. Au cours des années récentes, on a observé une variabilité croissante des salaires académiques tant pour les hommes que les femmes. Un écart entre hommes et femmes persiste qui est attribuable aux différences dans les rangs et dans les champs d'activités des hommes et des femmes. L'écart qui demeure inexpliqué a diminué dans le temps.

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1. Introduction

The Canadian academic labour market has undergone profound changes over the past 40 years. The professoriate has aged, and the percentage of faculty who are female has more than doubled. Collective bargaining, once relatively rare, is now standard, and all but a handful of Canadian universities are unionized (Chant 2006). Administrative discretion in salary setting has lessened as has, in many universities, the role of merit pay (Chant 2006). Equitable hiring and promotion practices, since the late 1980s, have been required by federal Employment Equity legislation,[1] while human rights and comparable worth legislation has guaranteed women – at least in theory – equal pay (McDonald and Thornton 2001). Government funding for universities has risen and fallen over this period with, as Martinello (2006) shows, direct impacts on academic salaries.

There is a rich Canadian literature exploring various aspects of the academic labour market: collective bargaining (Hum 1998; Hosios and Siow 2004; Chant 2006; Martinello 2009); provincial funding (Martinello 2006); merit pay (Chant 2006; Grant 1998); pay equity (McDonald and Thornton 2001); the status of female faculty (Ornstein, Stewart, and Drakich 1998), and the gender pay gap (Brown, Prentice, and Troutt 2007; Guppy 1989; Ornstein and Stewart 1996). While this literature illuminates broad trends, most authors have lacked either the data or the methods necessary to documents how *overall* trends impact *individual* academics.

In this paper we examine the labour market experiences of Canadian academics over the period 1970 to 2001. We begin by showing how entry earnings and returns to experience have varied across birth cohorts of male and female academics in Canada. Echoing the results of Beaudry and Green (2000) for the labour market as a whole, we find that each cohort of male academics in our sample has earned less than the previous one, with the exception of our youngest cohorts. Female earnings have been relatively stable across cohorts and, as a result, the male-female earnings gap has declined.

We next investigate the determinants of earnings with year-by-year cross-sectional analyses. Over time, the amount of the gender earnings gap that is 'unexplained,' or due to differences in the pay of women and men with the same education, age, rank, and field, has declined from more than 30% to less than 20%. This is consistent with a reduction in discrimination against female academics. Rank consistently explains more of the variation in earnings than any other factor, but is less important than it once was, while the returns to age and field have changed over time in interesting ways.

Finally, we use kernel density methods to estimate the male and female earnings distributions in 1970, 1980, 1990, and 2000, as well as a 'counter-factual density' – what the distribution of female earnings would have looked like had women been paid the same as men with the same observable characteristics. Over time, the

1 http://www.hrsdc.gc.ca/eng/lp/lo/lswe/we/information/history.shtml

spread of both the male and the female salary distribution has widened. Between 1970 and 1980, in particular, the overrepresentation of women at the bottom end of the salary distribution diminished, but men continue to be overrepresented at the top of the distribution, even when we control for observed characteristics.

We base our analysis on Statistics Canada data on the salaries of Canadian academics. This unique data set provides a census of all university teachers in Canada each year, allowing for both a national level of analysis as well as the inclusion of both personal- and institution-level information. While Guppy (1989) and Ornstein, Stewart, and Drakich (1998) use the same data set as we do, they analyze a different time period and employ different methodologies. Martinello's (2006, 2009) and McDonald and Thornton's (2001) studies are also based on the same data set; however, they use aggregated, institution-level data for Ontario universities. No previous analysis has been able to provide as comprehensive a picture of academic salaries in Canada as we do in this paper.

1.1. Data and sample selection

The data employed in the analysis are from the master files of the Full-Time University Teaching Staff Data of Statistics Canada over the period 1970 through 2001. This confidential, administrative database is collected every year by Statistics Canada from each of the universities in Canada. It contains detailed information on each employee's salary and rank, as well as personal information such as age, gender, and education. A particularly useful feature of the data is the inclusion of the name of the institution, allowing for analyses that include both institutional fixed effects as well as time-varying information at the university level.

The main measure of earnings employed is the actual salary earned by respondents.[2] This includes vacation pay, but stipends and honoraria for administrative duties are excluded.[3] We do not have information on other benefits or on teaching load, but we do know number of months of sabbatical and months on leave. We restrict our sample to people who did not have any unpaid leave in a given year.[4] The Consumer Price Index is used to convert nominal salaries into real ones.

Although the data has a longitudinal aspect, there exist challenges in terms of tracking individuals through time. It is not possible to follow the same individual from one university to another should he/she change jobs, and universities in a number of instances changed their individual identification variable over time (e.g., from an employee number to a Social Insurance Number or simply from

2 For medical/dental staff earnings include both the university contribution and the contribution by hospitals and medical care plans but excludes earnings from special grants and fees for services to patients.

3 Very similar results are found when we use annual rate of salary or total annual salary (annual rate of salary + administrative stipends).

4 The results are not sensitive to this restriction.

one individual identification number to an entirely new individual identification number). Therefore it is not possible to exploit the panel nature of the data.[5]

We include all full-time university teachers, including those who do not hold professorial ranks. Historically, discrimination against females in academia frequently took the form of appointment to instructor or lecturer positions (Ferber and Green 1982). More recently, US evidence suggests that substantial expansion in the growth of non-tenure-track faculty during the 1990s dampened growth in academic salaries; hence, non-tenure track faculty are an important part of the overall salary structure (Ehrenberg 2003, 269).

In table A1, the summary statistics for some key variables are presented for both men and women in 10-year intervals from 1970 to 2000. The average age of the sample increased over the period studied, especially for men, with the average age increasing from around 41 to 50 for men between 1970 and 2000. In 1970, female academics were on average a little less than two years older than males, but by 2000, male academics were around three years older than females. The proportion of females who are either full professors or associate professors also increased over the period studied, from around 28% in 1970 to around 61% in 2000. The proportion of faculty with PhDs also increased over this period, particularly the number of women. There are large gender differences in subject taught, although over time women have become less concentrated in the humanities, education, and nursing and have entered business/economics, the social sciences, and other disciplines.

2. Age-earnings profiles and period of labour market entry

The overall conditions present at the time of entry into the labour market are often thought to have lasting impacts on the age-earnings profile of a particular entry group. This is likely to be the case in the academic labour market, where tenure and collective bargaining mean that wage outcomes at the beginning of the employment relationship are likely to persist in terms of future earnings outcomes. Yet the start of our period, 1970, was characterized by administrative discretion in salary setting; collective bargaining became widespread during the 1970s and, to some extent, the 1980s.[6] During the middle of our period, the late 1980s and 1990s, pay equity legislation led Ontario universities, in particular, to adjust the salaries of female academics (McDonald and Thornton 2001). A cohort-based analysis is useful in exploring the extent to which these changes have altered the age-earnings profiles of individual generations of Canadian academics over time.

5 We examined different algorithms to recover individual identification numbers when individual faculty members left one university and joined another, or when universities changed their individual identification number but still had a large percentage that could not be consistently coded.

6 Information on collective bargaining provided by David Robinson of Canadian Association of University Teachers.

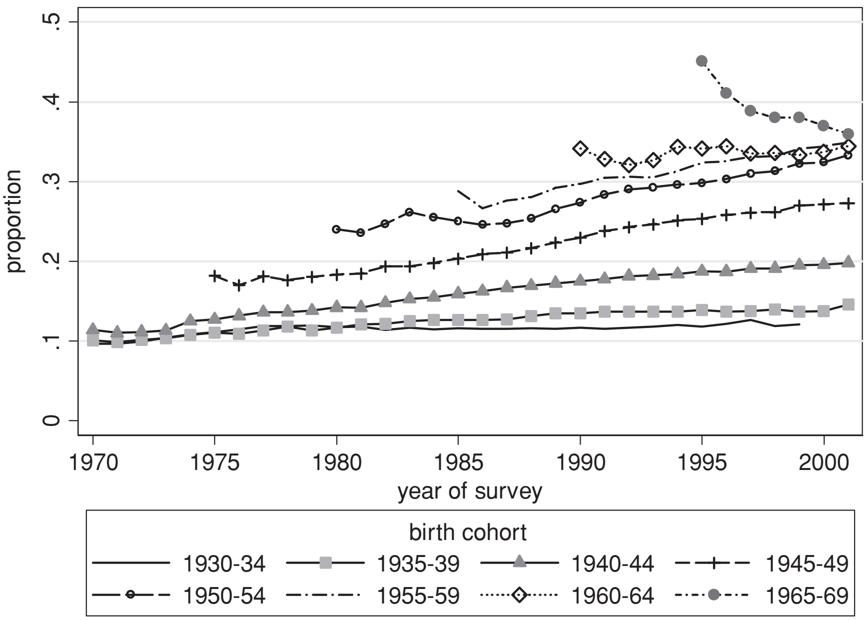

FIGURE 1 Proportion of university teachers who are females, by cohort
NOTES: Sample age 30–65 in the reference year, born between 1930 and 1969

The 1970 to 2001 period saw substantial numbers of women entering academia. Figure 1 shows the percentage of university faculty who are female by birth cohort.[7] For the earliest two cohorts, roughly 10% of faculty were women.[8] The percentage of university faculty who are female generally increased with every cohort until the present. For our middle cohorts, the proportion that were female rose over time but, by the 1960 to 1964 cohort, the pattern of latter female entry had disappeared.[9]

We estimated our cohort model of faculty salaries using the approach of Beaudry and Green (2000). The age-earnings profiles are estimated from regression models with the following specification:

7 We could instead carry out the analysis based on year of first entry into the labour market or year of first entry into the academic labour market. However, we do not have information on either variable; therefore, we define cohorts based on birth year.

8 For the cohort analysis, we restrict the sample to people aged 30 to 65 who were born between 1930 and 1969 and use five-year birth cohorts.

9 In 1995, only people born in 1965 and hired by age 30 are present in the 1965–69 cohort. This group had an unusually high proportion of females. Each year between 1995 and 2000 another birth year enters the 1965–69 cohort. For example, academics born in 1966 enter in 1996, and so on. Individuals hired after age 30 also join the sample. In 2000, when all 1965 to 1969 birth years are present, there is little difference between this cohort and other recent cohorts.

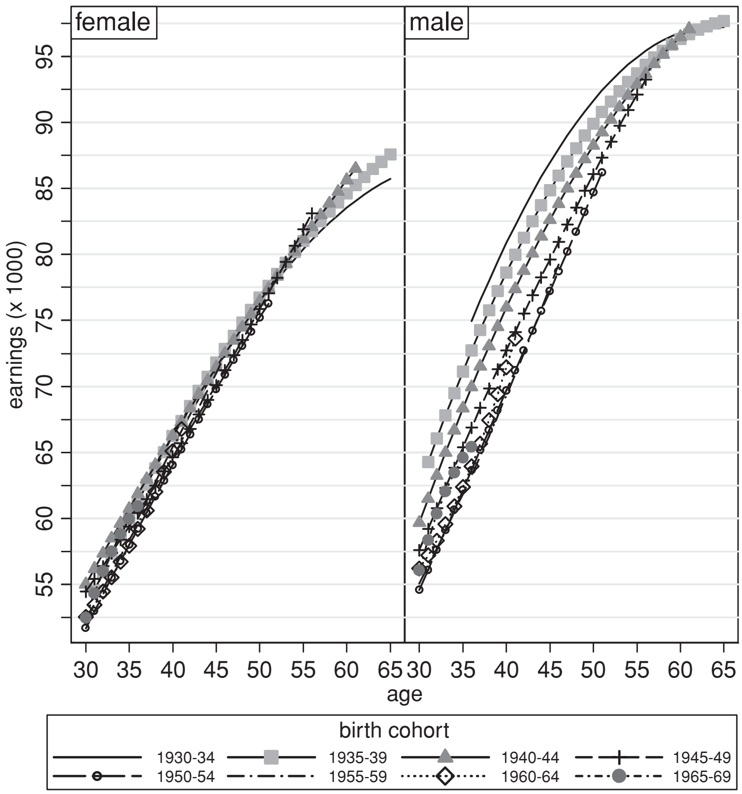

FIGURE 2 Age-earnings profiles

$$E_{it} = \sum_{j=1}^{J} C^j \left[\beta_0^j + \beta_1^j Age_{it} + \beta_2^j Age_{it}^2 + Male_{it}\left(\beta_3^j + \beta_4^j Age_{it}^2 + \beta_5^j Age_{it}^2\right)\right] + X_{it}\beta + \varepsilon_{it}, \tag{1}$$

where E_{it} is the earnings of individual i in year t. On the right-hand side are a set of dummy variables for the birth cohorts C^j for $j = 1, \ldots J$, that appear on their own and as interactions with both age (Age_{it}) and age-squared. $Male_{it}$ is an indicator variable that takes on a value of one for males. The specification in equation (1) allows initial salaries and earnings growth patterns to vary across birth cohorts for both the women and the men in our sample. In our estimates with 'full controls' we also include a vector, X_{it}, of controls for highest degree, place of first degree, place of highest degree, rank, subject taught, and, in some specifications, institutional fixed effects.

Figure 2 presents separate age-earnings profiles by gender, controlling only for age and age squared. There is no clear evidence of a shift in the age-earnings profile

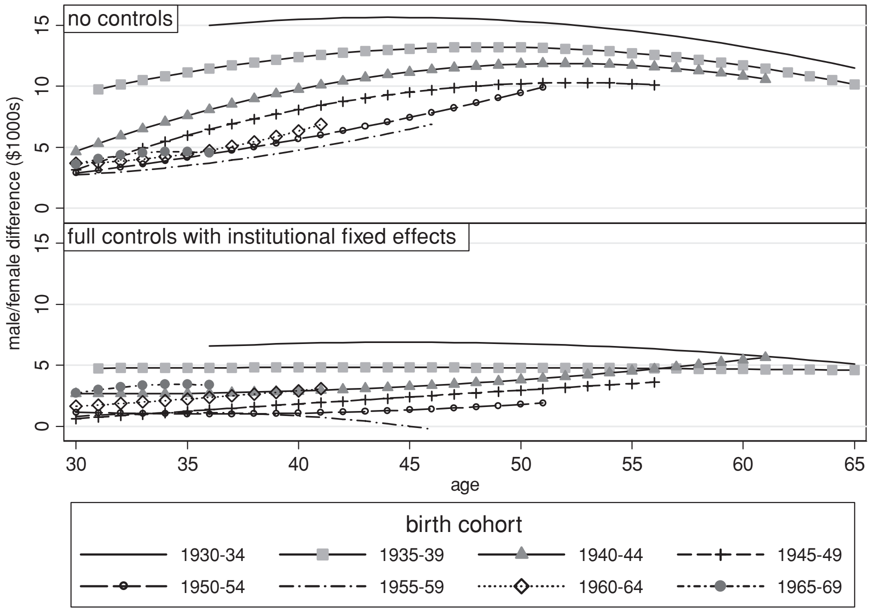

FIGURE 3 Male-female differences in earnings

of women across birth cohorts, in terms of either starting salaries or earnings growth.[10] Yet when we compare the age-earnings profiles between females and males, we see that women's earnings in academia have improved relative to those of men. Starting from the 1930–34 cohort, male earnings gradually fell in real terms through to the 1955–59 cohort. For the youngest cohorts, starting salaries have rebounded somewhat (the 1965–69 cohort's earnings curve is above the 1955–59 and 1960–64 cohorts' earnings curves).[11] However academics hired at a young age (the 30-year-olds at the far left-hand side of the earnings cohort) may be different from other academics, so it is hard to know if the 1965–69 cohort will continue to experience higher earnings. Figure 2 also shows that the rate of salary growth levels off for older individuals, especially men.

In the top half of figure 3, the differences between the male and female age-earnings profiles are presented by birth cohort (differencing the right-hand side and the left-hand side of figure 2). The male-female earnings differential has gradually narrowed across cohorts. This narrowing has partially taken the form

10 Restricting the age-earnings profiles to be equal for all cohorts, we find that the male cohort dummies are jointly statistically significant with an F-statistic of 317.08. As well, repeating the same test for females, we also find that the female cohort dummies are jointly statistically significant with an F-statistic of 6.58.

11 Using an F-test for separate intercept and age, age^2, we reject the equality between the 1965–69 and 1960–64 cohorts at the 1% level ($p = 0.0012$), and between the 1965–69 and 1950–54 cohorts also at the 1% level ($p = 0.0095$).

of better starting salaries for women – women in the baby boom cohorts, in particular, had a smaller earnings disadvantage at age 30 than the 1930s cohorts. The older cohorts experienced a decline in the raw male-female earnings differentials during the mid- to late-career years – the 40s, 50s, and 60s. By the 1950–54 cohort, the mid- to late-career narrowing of the gender gap had disappeared; however, we do not observe that cohort's late-career earnings. All of our cohorts, particularly the cohorts born after 1950, experience rising male-female salary differentials while in their 30s and 40s. The reasons why the gender earnings gap rises during the early- and mid-career years can be understood by comparing the profiles in the top half of figure 3 with those in the lower portion of the diagram.

In the bottom half of figure 3, we show the male-female earnings differentials across cohorts with all individual-level controls as well as institutional fixed effects (the picture without fixed effects; Warman, Woolley, and Worswick (WWW) 2006, figure 25, is similar). The absolute magnitude of the gender-salary differential once again declines across cohorts until the 1955–59 cohort, which, by mid-career, was experiencing gender equality. For our two youngest cohorts, however, gender-earnings differentials appear to be rising again.

Once we add all controls, there is no clear pattern to the change in the gender earnings gap as each cohort ages. The male-female earnings differential declined for the 1930–34 and 1955–59 cohort as those cohorts aged, remained static for the 1935–39 cohort and, for the remaining cohorts, increased slightly throughout their careers – but by substantially less than in the age-earnings profiles without controls. From this we can conclude that the widening of the male-female earnings difference during the early to mid-career years is primarily due to gender differences in our control variables: rank, institution, and discipline. In the next section, we show that rank explains a large portion of the gender earnings gap. Hence, our preferred explanation of the early- and mid-career increase in the gender earnings gap is differential promotion rates.

But why do many of our cohorts still experience some slight increase in the gender wage gap as they age even with all controls? Figure 1 shows that most of our cohorts became more female over time; the entry of modestly paid women or the exit of better-paid men would explain the patterns shown in figure 3. However, our controls for rank, institution, and discipline should pick up most of these composition effects. A separate cohort analysis of gender pay differentials at universities with and without merit pay (available in an on-line appendix linked to this article at http://economics.ca/cje/en/archive.php) found that the universities with merit pay generally had larger gender earnings gaps, as well as a more rapid widening of the earnings gap as each cohort aged. Hence, it seems likely that part of the gender earnings gap is due to men receiving higher discretionary salary increases, such as merit pay awards, or adjustments in recognition of outside offers.

It is interesting to interpret these findings in light of the general trends noted in the introduction to this paper. In economy-wide studies, collective bargaining has frequently been found to be associated with a reduction in the gender earnings gap (Blau and Kahn 2003). However, it does not appear, from our data, that the gender earnings gap experienced by older cohorts of female academics narrowed substantially once collective bargaining was introduced – the age earnings profile of the 1935–39 cohort, for example, is essentially flat once all controls have been introduced. The move away from administrative discretion and merit pay and towards seniority-based pay schemes instead entrenched existing inequalities (a finding confirmed by our analysis of seniority-based and merit-based universities in the on-line appendix). But while younger cohorts now experience something close to equality in starting salaries, the raw gender earnings gap widens as cohorts age. This begs the question: what determines academic salaries?

3. Determinants of academic salaries

In order to perform the cohort analysis in the previous section, we had to assume that the effects on earnings of rank, education, and subject taught did not change over time. In this section, we drop the focus on particular cohorts and instead explore how the determinants of academic salaries have evolved.

We estimate earnings models with a rich set of controls for the possible underlying determinants of academic salaries. The models are estimated separately for selected sample years in order to allow the coefficients on these variables to change through time. The following general form for the regression equations is employed:

$$E_i = \beta_0 + \beta_1 Male_i + \beta_2 Age_i + \beta_3 Age_i^2 + X_i\beta + \varepsilon_i \tag{2}$$

As before, earnings depend upon gender, age, and age squared.[12] The vector X_i, includes controls for highest degree, place of first degree, place of highest degree, rank, and subject taught. Rank measures productivity and experience (Ehrenberg 2003): we control for the usual professorial ranks (full, associate, assistant), as well as two ranks below assistant and an 'other' category. We classify scholars into 11 broad fields, such as 'humanities,' 'social sciences,' 'health' and 'engineering/applied science.' Some researchers have argued that the estimated coefficients on these variables are 'tainted,' in that discrimination against women may take the form of lower rates of promotion (Brown, Prentice, and Troutt 2007) or differential salaries across the academic disciplines. This possibility should be considered when interpreting the coefficients on these variables.

We measure human capital with highest degree and include age controls as proxies for years of work experience. We also include the country in which highest degree was awarded (Canada, US, UK, France, other), both as a measure of the quality of human capital and as a measure of initial bargaining position. A person with a US PhD, for example, can typically search over the entire North

12 Age is defined as age − 40.

American job market and hence is more likely to have an outside offer when negotiating his/her starting salary. We control for country in which first degree was awarded because we believe this could have an effect through the longer time-to-completion of immigrants or the effect of restrictions on the hiring of non-Canadians to academic positions. We experimented with including additional controls based on institutional information (e.g., a 'francophone' control for all French-language universities). While these controls were frequently significant and of independent interest, they did not completely control for the multiple differences between institutions that influence salaries (local cost of living, weather). We instead controlled for salary variation coming from institutional differences through institutional fixed effects.

Table 1 shows the results of regression analysis of academic earnings in 1970, the start of our sample period, and again in 2000. We note major differences that occur in the other years of the sample or if the 1970 and 2000 comparison does not follow the general trend. The base case for the regression analysis is a female, 40 years of age, who is a full professor with a PhD, who took all of her degrees in Canada, and teaches in the social sciences. In the first two columns, the coefficient on the $Male_i$ dummy gives the average difference between male and female academics' salaries in 1970 and 2000 (in 2001 dollars) with no control for any differences between male and female academics. The next two columns provide estimates from OLS regression specifications containing a broad set of control variables but without institutional fixed effects; columns 5 and 6 provide OLS estimates from the same regression specifications but with institutional fixed effects included.

Comparing the 1970 and 2000 results, we find that the raw salary differential has fallen by just over 30%. The male wage premium including all individual controls has fallen by more than 60%; with all individual controls and institutional fixed effects, the male wage premium has fallen by just over 50%.

The coefficient on the age-square term switches from being negative in the 1970 results to positive in the 2000 results in table 1. Normally, one would expect diminishing returns to work experience to lead to a negative coefficient on the age-square term. However, these cross-sectional profiles combine both the returns to experience within cohorts and differences in earnings across cohorts. Because older academics had much higher salaries relative to younger academics in 2000 compared with 1970,[13] we see an increase in the return to age in the cross-section.

The returns to rank have changed substantially over time: the salary gap between a full professor and a lower-ranked academic declined by over $10,000 between 1970 and 2000, with academics at the lower ranks making the greatest gains.[14] Place of highest degree, which we hypothesized was a measure of degree quality and negotiating position, has a significant effect on salary. Between 1970

13 We reject the equality of the age slopes between 1970 and 2000 at the 1% level.

14 We reject both the individual and joint equality of the 1970 and 2000 rank coefficients at the 1% level.

TABLE 1
Determinants of real earning: regression analysis

	1970	2000	1970	2000	1970	2000
Male	16,785*	11,635*	5,095*	1,977*	5,068*	2,397*
	[936]	[654]	[467]	[299]	[537]	[247]
Age			434.74*	459.40*	461.84*	501.49*
			[32.57]	[63.27]	[30.83]	[55.47]
Age^2			−9.97*	11.10*	−10.22*	8.14*
			[1.69]	[2.81]	[1.64]	[2.38]
Rank (default full professor)						
Associate			−25,585*	−14,329*	−25,269*	−13,884*
			[1,071]	[1,224]	[1,031]	[1,056]
Assistant			−37,873*	−24,319*	−37,112*	−24,203*
			[1,303]	[1,509]	[1,262]	[1,440]
Below assistant			−47,415*	−31,513*	−46,698*	−32,223*
			[1,506]	[2,327]	[1,454]	[1,620]
Below below assistant			−51,282*	−28,073*	−51,366*	−32,223*
			[2,840]	[1,964]	[3,358]	[2,146]
Other			−40,307*	−21,506*	−40,223*	−19,600*
			[2,892]	[2,756]	[3,038]	[2,786]
Highest degree (default PhD)						
Professional			6,542+	5,979	7,483*	3,363
			[2,519]	[6,068]	[2,630]	[6,896]
Graduate			−2,336*	−1,736*	−1,267*	192
			[407]	[495]	[270]	[561]
Undergrad			−1,004	−1,205	−967^	176
			[889]	[1,432]	[551]	[1,291]
Other degree			−7,481*	272	−6,526*	−3,871+
			[1,252]	[1,422]	[1,069]	[1,824]
Place of highest degree (default Canada)						
Other country			1,023	1,132	658	1,276^
			[775]	[728]	[687]	[673]
US			670	3,836*	396	2,171*
			[540]	[643]	[507]	[353]
UK			1,656*	2,516*	1,462*	1,827*
			[412]	[468]	[266]	[392]
France			−1,199	−640	637	788^
			[809]	[677]	[748]	[398]
Place of first degree (default Canada)						
Other country			−1,756*	−1,173*	−2,262*	−591+
			[620]	[386]	[485]	[277]
US			278	201	−314	−296
			[409]	[792]	[312]	[291]
UK			−148	1,570^	−906*	606
			[477]	[834]	[324]	[581]
France			−151	−183	−657	−87
			[685]	[635]	[676]	[503]
Main subject taught (default social sciences)						
Education			3,086^	−561	182	−252
			[1,688]	[796]	[595]	[562]
Fine Arts			−2,464*	−4,439*	−2,712*	−4,205*
			[861]	[791]	[679]	[618]
Humanities			−2,842*	−3,674*	−2,256*	−2,613*
			[446]	[634]	[363]	[472]

Continued

TABLE 1 *Continued*

	1970	2000	1970	2000	1970	2000
Business\economics			2,001*	7,867*	2,319*	7,820*
			[712]	[1,128]	[531]	[1,178]
Engineering\appl scien			651	4,921*	451	3,185*
			[971]	[793]	[723]	[518]
Agricultural\biol science			−1,128	607	−1,005^	328
			[923]	[857]	[596]	[523]
Nursing			−3,922*	237	−4,391*	866
			[890]	[1,403]	[768]	[1,078]
Health			11,959*	5,595*	11,673*	4,358+
			[1,143]	[1,906]	[1,144]	[1,690]
MathPhySci			54	1,808*	307	1,594*
			[573]	[561]	[449]	[348]
Other subject			10,086*	−5,171^	8,783*	−1,303
			[1,881]	[2,930]	[2,458]	[2,175]
Constant	62,902*	72,357*	98,812*	82,031*	94,043*	89,831*
	[752]	[1,218]	[1,630]	[1,714]	[1,138]	[1,321]
Observations	20,089	32,815	20,089	32,815	20,089	32,815
R-squared	0.06	0.05	0.73	0.46	0.76	0.56
With institutional fixed effects	NO	NO	NO	NO	YES	YES

NOTES: Robust standard errors in brackets are clustered on institution. Values are in 2001 Canadian dollars. ^significant at 10%; +significant at 5%; *significant at 1%.

and 2000, the premium associated with having an American highest degree rose substantially, while those with highest degrees from the UK continued to receive a significant salary premium relative to Canadian-educated academics.[15]

The results in table 1 also show how the earnings penalty or premium associated with different disciplines has changed over time relative to the default category of the social sciences. The traditionally female dominated discipline of nursing has seen a relative rise in salaries. In contrast, the health category is rewarded less generously than it once was. The wage premium associated with business and engineering has risen over time, while the earnings disadvantage for those in the fine arts and humanities has increased in absolute value.[16] Except in the humanities, engineering/applied science, and health, the size of the coefficients on the various discipline controls does not change substantively when institutional fixed effects are added, suggesting that they arise from factors outside particular

15 We reject the equality of the 1970 and 2000 US highest degree coefficients at the 1% level, but we cannot reject the equality of the 1970 and 2000 UK highest degree coefficients at conventional levels.

16 The disadvantage in fine arts increases between 1970 and 1971 but is fairly stable thereafter. The change in the premium or penalty relative to social sciences between 1970 and 2000 for Nursing, Business/Economics, Engineering, and Health are all statistically at the 1% level, whether or not institutional fixed effects are included. For fine arts, the increase in the penalty is statistically significant at the 1% level without institutional fixed effects, but at only the 5% level with fixed effects.

university settings, such as market conditions. When we rerun the results using log earnings as the dependent variable, we obtain very similar findings, in terms of both the male-female earning differential, as well as the change in the returns to the other variables over time.[17]

Next we use the Blinder (1973) / Oaxaca (1973) decomposition to break down the average gender earnings differential, $\bar{Y}_M - \bar{Y}_F$, into explained and unexplained portions, as shown below in equation (3). The 'explained portion' is the difference in average male and female characteristics, $\bar{X}_M - \bar{X}_F$, evaluated at β^*, which we calculate as $(\beta_M + \beta_F)/2$. Here, $\hat{\beta}_M$ and $\hat{\beta}_F$ come from estimating equation (2) separately for males and females, respectively.[18] The unexplained portion, the final term in equation (3), can be interpreted as differences in the returns that men and women receive from any given characteristic.

$$(\bar{Y}_M - \bar{Y}_F) = \underbrace{(\bar{X}_M - \bar{X}_F)'\beta^*}_{Explained} + \underbrace{\left[\bar{X}'_M(\hat{\beta}_M - \beta^*) + \bar{X}'_F(\beta^* - \hat{\beta}_F)\right]}_{Unexplained} \tag{3}$$

Table 2 shows the results of the Blinder-Oaxaca decomposition. We show the total 'explained' and 'unexplained' earnings gap and also break the explained and unexplained portions of our decomposition into subcategories. Given the sensitivity of the contribution of a group of categorical variables to the choice of the omitted category (Oaxaca and Ransom 1999), we use deviation contrast transformations of the categorical variables (Yun 2005).[19] Table 3 provides the background regressions. Because of space constraints, we report only the results with institutional fixed effects.

The fraction of the earnings differential that can be explained is just under 70% in 1970 and around 80% in 2000, although it is likely that differences in observable factors that we could not control, for example, the type of research published or the allocation of time between teaching and research, account for further differences. The upward trend in the percentage 'explained' is most pronounced between 1970 and 1990 (for year by year decompositions, see WWW 2006, figures 27, 28).

At the start of the period, differences in rank accounted for almost 90% of the earnings differential attributable to differences in characteristics shown in table 2 and more than 60% of the overall difference. Throughout our period, as women entered the professorial ranks and as the returns to rank fell, it became less

17 For example, the unconditional earnings difference is 26% in 1970 and 16% in 2000, while with controls it is around 7% in 1970 and 3% in 2000 (regardless of whether institutional fixed effects are included).

18 The results were re-estimated with $\beta^* = \beta_F$, as well as $\beta^* = \beta_M$. There was a slight difference in magnitude, but the overall interpretation does not change.

19 Also see Nielsen (2000) and Gardeazabal and Ugidos (2005) for other solutions to the identification problem. In view of the fact that the contribution of age to the differences in returns to characteristics is very sensitive to the type of transformation, we omit both the age and the constant from the differences in the returns to characteristics in table 2 (see Oaxaca and Ransom 1999 and Yun 2005).

TABLE 2
Oaxaca decomposition results

	1970	2000	1970	2000
Total male/female difference	**16,785**	**11,636**	**16,785**	**11,636**
Attributable to differences in characteristics	**11,531***	**9,589***	**11,539***	**9,209***
	[936]	[623]	[1,352]	[1,868]
Age	−407*	1,886*	−454*	1,858*
	[120]	[183]	[131]	[509]
Rank	10,296*	5,799*	10,128*	5,869*
	[689]	[513]	[680]	[892]
Highest degree	742*	270	563*	55
	[155]	[183]	[134]	[1,798]
Place of last degree	42	256*	34	177*
	[109]	[278]	[91]	[228]
Place of first degree	39	31	24	10
	[144]	[336]	[161]	[33]
Subject	818+	1,349*	846+	963*
	[375]	[232]	[347]	[176]
Institutional fixed effects			398	277
			[792]	[1,724]
Attributable to differences in returns to characteristics	**5,254***	**2,047**	**5,246***	**2,427***
	[849]	[1,600]	[517]	[398]
Rank	−1,327	−1,171	−1,087	−1,310
	[949]	[907]	[963]	[526]
Highest degree	−1,227^	577	−1,532*	1,038
	[649]	[1,217]	[562]	[165]
Place of last degree	−126	78	−354	−64
	[248]	[83]	[237]	[47]
Place of first degree	165	20	321	−129
	[232]	[50]	[234]	[275]
Subject	−826*	79	−867*	53
	[301]	[405]	[313]	[353]
Institutional fixed effects			3,091*	2,303*
			[378]	[716]
Observations	20,089	32,815	20,089	32,815
With institutional fixed effects	NO	NO	YES	YES

NOTES: Robust standard errors in brackets are clustered on institution. Values are in 2001 Canadian dollars. ^ significant at 10%; + significant at 5%; * significant at 1%.

important as a source of explained variation, accounting for around 60% of the explained gender earnings differential in 2000. Men's and women's returns to rank also became more similar over time. As table 3 shows, in 1970, the gap between associate and full professors was smaller for women than for men, while by 2000, the associate/full salary gap was between \$13,500 and \$14,000 for both women and men (with all controls).[20] Yet the fact that women hold lower ranks than men still accounted, in 2000, for about 50% of the overall earnings differential.

20 In 1970, the difference is statistically significant at the 1% level, while in 2000 the difference is not statistically significant.

TABLE 3
Regressions by gender, with institutional fixed effects

	Male 1970	Female 1970	Male 2000	Female 2000
Age	517.33*	167.46*	500.68*	524.27*
	[35.95]	[26.00]	[73.28]	[52.17]
Age^2	−11.31*	−2.37	8.99*	3.98
	[1.88]	[1.97]	[3.13]	[2.42]
Rank (default full professor)				
Associate	−25,086*	−20,394*	−13,929*	−13,492*
	[1,037]	[852]	[1,179]	[857]
Assistant	−36,592*	−34,528*	−24,406*	−23,763*
	[1,298]	[1,033]	[1,626]	[1,179]
Below assistant	−45,576*	−45,247*	−32,356*	−31,942*
	[1,493]	[1,099]	[2,150]	[1,435]
Below below assistant	−49,229*	−50,493*	−30,970*	−32,146*
	[3,582]	[1,780]	[2,237]	[2,183]
Other	−39,464*	−39,344*	−16,944*	−23,884*
	[3,147]	[4,072]	[3,105]	[2,198]
Highest degree (default PhD)				
Professional	7,914*	561	2,336	6,078
	[2,745]	[2,182]	[6,592]	[8,152]
Graduate	−1,449*	−1,218+	409	−6
	[306]	[586]	[628]	[561]
Undergrad	−722	−2,882*	167	137
	[544]	[850]	[1,409]	[1,451]
Other degree	−5,087*	−6,536*	−4,668+	−1,900
	[1,056]	[1,068]	[2,268]	[1,390]
Place of highest degree (default Canada)				
Other country	674	1,373	1,646+	402
	[758]	[833]	[782]	[974]
US	448	598	2,309*	1,618*
	[508]	[666]	[407]	[377]
UK	1,639*	−179	1,755*	1,885*
	[281]	[800]	[416]	[671]
France	754	−264	763^	1,197+
	[773]	[1,104]	[431]	[597]
Place of first degree (default Canada)				
Other country	−2,570*	−1,269+	−670^	−578
	[483]	[628]	[348]	[453]
US	−433	267	−679+	761^
	[332]	[592]	[339]	[457]
UK	−1,147*	906	407	1,187
	[332]	[611]	[646]	[750]
France	−795	−231	630	−1,369^
	[757]	[942]	[636]	[813]
Main subject taught (default social sciences)				
Education	−102	1,421	−355	102
	[581]	[1,052]	[644]	[567]
Fine Arts	−3,000*	−2,401+	−4,444*	−3,758*
	[741]	[941]	[703]	[684]
Humanities	−2,467*	−1,570*	−2,765*	−2,205*
	[407]	[578]	[555]	[490]
Business\economics	2,679*	−1,394	7,396*	8,674*
	[520]	[1,250]	[1,222]	[1,279]

Continued

TABLE 3 *Continued*

	Male 1970	Female 1970	Male 2000	Female 2000
Engineering\appl scien	587	−1,500	2,886*	3,431*
	[764]	[3,217]	[506]	[899]
Agricultural\biol science	−1,037	−802	102	799
	[648]	[900]	[558]	[650]
Nursing	−5,711+	−3,397*	−668	1,780^
	[2,242]	[602]	[2,330]	[965]
Health	12,358*	4,953*	4,240+	4,423*
	[1,154]	[1,323]	[1,948]	[1,232]
MathPhySci	499	−173	1,297*	2,074*
	[463]	[940]	[393]	[608]
Other subject	9,861*	68	−846	−1,935
	[2,410]	[1,308]	[2,083]	[3,119]
Constant	97,888*	95,454*	93,249*	83,706*
	[1,083]	[1,157]	[1,573]	[1,184]
Observations	17,700	2,389	23,624	9,191
R-squared	0.75	0.78	0.53	0.55

NOTES: Robust standard errors in brackets are clustered on institution. Values are in 2001 Canadian dollars. ^significant at 10%; +significant at 5%; *significant at 1%.

In 2000, the average female academic was younger than the average male academic, and older academics earned considerably more than younger academics. Taken together, this meant that age accounted for about 20% of the explained earnings gap and 15% of the overall difference in earnings in 2000, as shown in table 2. Interestingly, however, age differences contributed *negatively* to the explanation of the male-female earnings gap at the beginning of our period.[21] The freshly minted PhDs hired during the rapid university expansion (and hot job market) of the 1960s were predominantly male. The women in academia in 1970 were older than the men, and hired during a time when the rewards to being an academic were lower (Johnson 1966 provides historical data on academic salaries in Canada).

As noted earlier, it can be argued that inclusion of rank as an explanatory variable is problematic because rank itself may reflect discriminatory promotion practices. When we rerun the results without controlling for rank (results not shown), the amount of the earning difference attributable to differences in characteristics is only around $3,500 and $6,700 in 1970 and 2000, respectively, and around $3,800 and $6,300 when we add institutional fixed effects. Without controlling for rank, the negative effect of age is even larger in 1970 and the positive effect is higher in 2000.

Tables 2 and 3 also show how disciplinary penalties and premia have evolved over time. Overall, subject taught is a significant source of explained variation in 2000, but not in 1970 (see table 2). In 1970, there was not an obvious advantage

21 Age becomes positive in the late 1970s.

TABLE 4
The determinants of earnings at medical, comprehensive, and undergraduate universities

	1970			2000		
	Medical	Comp	Undergrad	Medical	Comp	Undergrad
Male (no controls)	*18,946**	*14,846**	*10,010**	*12,183**	*10,716**	*10,242**
	[1,137]	*[1,855]*	*[1,159]*	*[955]*	*[1,144]*	*[1,322]*
Male	5,785*	3,026*	3,228*	2,032*	2,057*	2,530*
	[578]	[571]	[970]	[401]	[541]	[400]
Associate	−26,506*	−24,791*	−19,608*	−16,453*	−10,259*	−13,120*
	[1,394]	[1,739]	[1,179]	[1,700]	[1,045]	[1,069]
Assistant	−38,283*	−37,622*	−33,079*	−26,132*	−19,624*	−22,779*
	[1,767]	[2,092]	[1,217]	[2,362]	[1,447]	[1,514]
Other controls included but not reported: highest degree, place of first and highest degree, ranks below assistant professor, age age-squared, and a set of 10 field identifiers						
Observations	12,430	4,513	1,701	17,932	9,588	4,260
R-squared	0.71	0.8	0.81	0.4	0.59	0.66

NOTES: Robust standard errors in brackets that are clustered on institution.
^significant at 10%; +significant at 5%; *significant at 1%

for women to being in a more male-dominated discipline.[22] Women in economics and business, engineering, and math and physical sciences may have earned slight wage penalties relative to those in the social sciences.[23] By 2000, however, women in these disciplines earned at least as great a premium as men in these disciplines. Women and men for the most part experience the same discipline-specific rewards. In 1970, that was not always the case; for example, men in health earned $12,358 more than men in the social sciences, while women earned just $4,953 more.[24]

By 2000, there was not much unexplained variation in men's and women's earnings. Interestingly, the major source of unexplained variation in table 2 is 'institutional fixed effects,' which accounted for almost 60% of the unexplained variation in 1970, and about 95% of the unexplained gender earnings differential in 2000. In other words, the unexplained salary gap largely stems from the fact that some institutions pay men more than women. From the analysis shown in table 2, however, we do not know which institutions or why.

To explore this issue further, we reran all of the results separately for different types of universities and examined how the relationships between earnings and personal characteristics varied across three broad categories of universities: (1) universities with medical schools, (2) comprehensive universities, and (3) primarily undergraduate universities. The results are presented in table 4. The universities are classified according to their status in 2000, based on the *Maclean's Magazine*

22 The male subject coefficients are jointly significant at the 1% level every year from 1970 to 1988, when we allow all coefficients to vary by gender. For the remaining years, the level of significance varies and in several years we cannot reject that the disciplinary penalties or premia are equivalent.
23 1970 is the only year where men earn a large premium in business and economics.
24 Males experience a statistically significant earning premium in health every year until 1992.

categorization of universities. This analysis exposes some interesting findings. In 1970, universities with medical schools had the greatest gender gap, based on analyses both with and without controls, but by 2000, there were only very small differences in the gender gap by type of university.[25] While the returns to rank declined at all schools, the greatest decline was experienced at comprehensive schools. In the 1970s, the return to being a full professor relative to either an assistant or an associate professor was similar at comprehensive and medical schools, but starting in the 1980s, the relative advantage enjoyed by full professors at comprehensive schools declined by around $4,000.[26] For most years, the advantage experienced by full professors was around $3,000 or $4,000 less at primarily undergraduate schools than it was at universities with medical schools.

For the most part, the disciplinary premia did not show a consistent pattern across these types of universities. One exception is business/economics, where the universities with medical schools were able to pay much higher salary premiums than the comprehensive and undergraduate universities.[27] As well, while the premium for business/economic professors increased over the period studied at all three types of schools, the growth in the premium was largest at universities with medical schools.[28] Another exception is the penalty to being in fine arts, which is usually smaller at universities with medical schools.

In the labour market as a whole, the gender earnings gap is most pronounced at the top end of the earnings distribution: there are smaller differences between the earnings of lower-paid men and women (Blau and Kahn 2000). Our findings are suggestive of a similar pattern in the academic labour market. Universities with medical schools pay certain professors in certain disciplines, such as medicine and business/economics, very well.[29] To the extent that men are more likely to be found in these highly paid positions, this phenomenon at least partly explains the gender pay gap.

In order to compare our results with those of other studies, we reran the regressions in tables 1 and 3 with log earnings as the dependent variable (the results of this exercise are available from the authors). Using same-year (1999) data, we

25 During the 1990s, the male earnings advantage at primarily undergraduate schools is slightly larger by around $1,000 or $1,500 relative to comprehensive and medical schools, respectively.

26 In the 1970s, we cannot reject the equality of either the associate or assistant professor coefficients for comprehensive and medical schools. Starting in the early 1980s, the difference in the relative gap between associate and full professors is statistically significant at either the 1% or the 5% level, while the difference in relative gap for assistant professors is not statistically significant in most years.

27 The premium is actually larger at comprehensive schools from the mid-1970s until the mid-1980s, although the coefficient on the difference in the premium between comprehensive and medical schools is not statistically significant.

28 In 2000, business/economic faculty experienced a $11,000 earnings advantage over social science faculty at medical schools; this advantage was around $5,000 lower at comprehensive and undergraduate schools (the difference between the business/economic coefficients at medical school and comprehensive and undergraduate schools is statistically significant at the 5% and 10% levels, respectively). As well, in most years, the business/economic earnings advantage is smaller at undergraduate schools relative to comprehensive schools.

29 Individual salaries are documented on the Ontario Salary Disclosure Web site, http://www.fin.gov.on.ca/english/publications/salarydisclosure/2008/univer08.html.

TABLE 5
Comparison of Canadian and US gender-earnings differentials: coefficient on male dummy variable

	Basic controls and field (no rank)		Basic controls and rank (no field)	
	Barbezat	WWW	Barbezat	WWW
1968/70	0.165 (0.015)	0.146 (0.009)	0.151 (0.015)	0.068 (0.007)
1975	0.104 (0.013)	0.134 (0.009)	0.099 (0.013)	0.076 (0.006)
1977	0.046 (0.009)	0.107 (0.008)	0.049 (0.009)	0.059 (0.004)
1984	0.068 (0.011)	0.087 (0.006)	0.068 (0.011)	0.051 (0.004)
1989	0.070 (0.009)	0.075 (0.005)	0.107 (0.010)	0.046 (0.004)

NOTES: Standard errors in parentheses. Canadian error terms are robust standard errors clustered on institution. US error terms are computed in the standard way.

found gender-earnings gaps comparable to those reported in Blackaby, Booth, and Frank's (2005, F87) study of UK economists: with controls for only male and ethnicity, they found a 19% earnings premium for being male (coefficient of 0.177 and using $e^{.177} - 1$), whereas we found a 16% earnings advantage for all male academics in a regression controlled only for gender. In their most comprehensive regression (which includes productivity and personal characteristics as well as institutional characteristics) they still found a 10% earnings advantage for males, whereas we found only a 3% earnings advantage with all controls for the same year.

Table 5 compares our results with Barbezat's (1991, 193) US study. To make our results comparable with hers, we restricted our sample to assistant, associate, and full professors and added in provincial dummies to control for region. However, the regression specifications are not identical – Barbezat controls for time spent in administration and teaching versus research. Without rank controls, we found, as she did, a reduction in the gender-earnings differential, particularly between 1968/70 and 1977. The gender earnings differentials without rank are similar in the two countries (except in 1977), although the Canadian differential is slightly larger. The similarity of the gender-earnings differentials in Canada and the US at the beginning and end of our period is consistent with the idea that academic salaries are determined within a North American labour market – although the smaller size of the Canadian market might be expected to create more possibilities for market power, especially in the francophone job market. With controls for rank, we generally found a smaller gender difference in Canada. Table 2 shows that rank accounts for a large part of the gender-earnings differential in Canada; the findings in table 5, showing a smaller earnings differential when rank is controlled for, are another aspect of the same phenomenon.

4. Comparing the earnings distributions of male and female faculty

Regression analyses effectively portray the evolution of, and determinants of, average salaries. Yet one of the stylized facts in the US literature is that the variance of the academic salary distribution is increasing (Ehrenberg 2003). In this section,

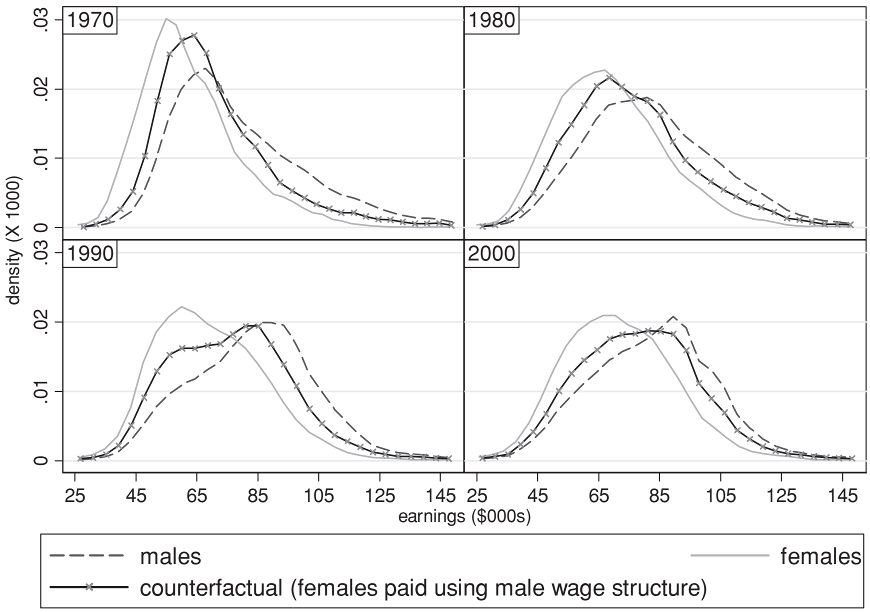

FIGURE 4 Density of earnings and counter-factual estimates, by selected year

we document the evolution of three salary distributions over time: men's, women's, and a female 'counter-factual' salary distribution, following the approach suggested by DiNardo, Fortin, and Lemieux (1996). The 'counter-factual' salary distribution shows, given observable characteristics, what women would have earned had they been paid based on the male wage structure. These estimates are created by re-weighting the males' earnings function, taking into consideration the observable characteristics of females, to generate a counter-factual distribution, which can then be estimated using the kernel density estimation procedure.

Figure 4 shows the distribution of male, female, and counter-factual academic salaries in 10-year intervals from 1970 to 2000. This figure is generated from separate kernel density estimates of the male and female salary distributions,[30] with salary data normalized to 2001 Canadian dollars. As in the US, both male and female salary distributions have widened, particularly between 1970 and 1980.[31] Part of this widening reflects demographic change. The rapid university

30 When we use the Kolmogorov-Smirnov test, equality of the female and male distributions is rejected in each year at the 1% level.

31 We also reran these estimates using the same restriction as those used in the cohort analysis (born between 1930 and 1969). We found that this effect is much more pronounced since as the years of the sample progress, the earlier cohorts are aging while newer cohorts are entering which creates more variance in age in the later years, and consequently we see more variance in earnings.

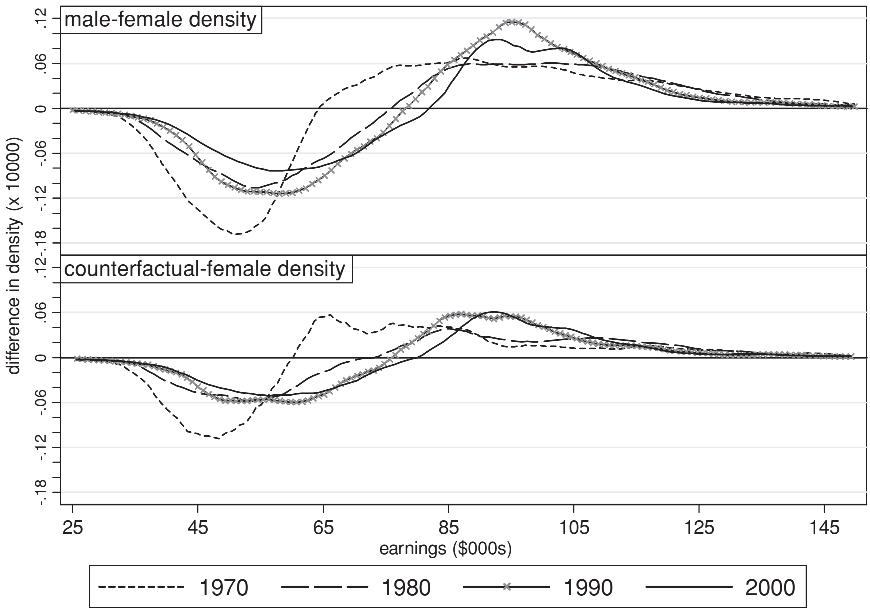

FIGURE 5 Difference in male/female density and counter-factual/female density

expansion of the 1960s created a relatively compressed salary structure. The peak of the male salary distribution, in particular, moves to the right over the decades following the large 1960s-expansion academic bulge. At the same time, the increased spread of the salary distribution is consistent with the story told in the previous section: at universities with medical schools, the returns to economics/business are increasing, and relative to other universities, the gap between the full professors and assistants/associates is widening (see table 4). Throughout our period, the salary structure for women is more compressed than the salary structure for men, at first because of the concentration of women in the lower ranks and then (in the latter part of our sample period) because of the concentration of women at the younger age levels.

In the top half of figure 5, the four pairs of male and female salary distributions from figure 4 are differenced. Values above zero represent the range of the distribution where the male density is greater than that of females. Points below zero represent the range where the density of females is greater. The dip down in this figure shows females are relatively overrepresented at earnings of $50,000 to $60,000; the rise upward indicates the overrepresentation of men at the $90,000 to $100,000 range of the earnings distribution. Over time, the bottom of the 'dip' has shifted from $50,000 to $60,000, as the modes of both men's and women's salary distributions have increased. The downward dips in figure 5 have also flattened out over time, as the earnings distributions for women and (and to a lesser extent men) have become markedly less concentrated. The greatest conversion in salaries happened between 1970 and 1980. A separate analysis by age (see WWW 2006, figures 9e, f) shows that the greatest 'flattening out' of the downward dip has occurred in the 30–39 age group as starting salaries have converged. The pattern of male overrepresentation – the peak in the top half of figure 5 – has also changed over time. Rather than flattening out, male overrepresentation at around $100,000 became more pronounced between 1970 and 1990. Essentially, the peak of the earnings distribution tracks the large group hired in the 1960s and shifts to the right as their salaries increase.

The counter-factual density in figure 4 shows what the distribution of female earnings would have been had women been paid according to the male salary function. In every decade, the counter-factual distribution is to the left of the male distribution. Women would have earned more had they been paid based on the male salary structure, and the salary distribution would have been less compressed. The gap between the male and the counter-factual density functions (not plotted) represents the gender earnings gap due to gender differences in the observable characteristics discussed in the previous section: women are in the lower ranks and differ from men in their fields and qualifications.

The gap between the female and the counter-factual distribution, plotted in the lower portion of figure 5, is the two-dimensional equivalent of the 'unexplained' earnings differential in table 2. This gap represents the difference between how the university labour market rewards men and women with the same characteristics *at each part of the earnings distribution*. Points below zero represent ranges of the earnings distribution where there are *more* women than one would expect based on observable characteristics; points *above* are ranges where women are *under-represented*. The difference between the female and counter-factual distributions shrank between 1970 and 1980; since 1980 it is difficult to detect further convergence. A separate analysis by age (WWW 2006, figure 36a2) shows convergence in the salaries of 30–39-year-old academics, particularly between 1970 and 1980. Women between 50 and 59 (WWW 2006, figure 36a3), since 1980, have begun to be underrepresented at earnings just under $100,000. The overall story is similar to that told by the cohort analyses: convergence in starting salaries between the Depression-era and baby-boom cohorts, gender-earnings salary gaps that widen with age for the more recent cohorts.

5. Conclusions

New PhD graduates negotiating their first employment contracts at universities are often advised that 'the intercept is negotiable, but the slope is fixed.' To some extent, our paper supports this conventional wisdom. The cohort analysis in section 2 demonstrates just how much starting salaries matter. Males in the cohorts born prior to 1945, many of whom were hired during the rapid university expansion of the 1960s, have enjoyed higher salaries than members of any other cohort throughout their entire academic careers. This fact is even more striking given the institutional changes that have taken place in universities during the period of our analysis. Despite widespread introduction of collective bargaining, pay equity, and employment equity legislation, the women in our older cohorts were paid less than their male contemporaries throughout their working life.

This shows how difficult it is to eliminate existing earnings differentials. A single instance of discrimination during initial salary negotiations can have effects that last for decades. At the same time, this suggests some important general lessons from our analysis. First, it is crucial to ensure equity at the time of hiring. Second, any evaluation of progress towards gender equity within academia has to take into account institutional inertia. We would like to know more about this inertia and why institutions did not do more to remedy the earnings disadvantages experienced by our older female cohorts.

The gender-earnings gap continues to rise with years of experience for each generation, peaking in the mid-career years of the 40s and 50s. Even women who start off in a position of equality with their male contemporaries will fall behind over time if they are promoted more slowly or receive smaller salary increments. Furthermore, initial job market conditions influence a cohort's entire lifetime earnings profile. Men who entered the difficult academic labour market of the 1980s and 1990s – our 1955–59 and 1960–64 cohorts – fared worse than members of earlier generations and, it appears, later ones. Today men and women are part of the same academic labour market and are subject to the same market conditions.

Cross-sectional estimates support the idea that the male and female salary structures are converging. The earnings differential that can be attributed to differences in women's and men's rank, age, education, and field is, by the end of our study period, around 80%. While rank is still the most important explanation of the male-female earnings gap, age has now emerged as a major contributor, reflecting differences in experiences across cohorts. In many cases, salary differentials across fields have widened over our period of study, but, by 2000, no significant part of the gender earnings differential could be attributed to systematic differences in the way men and women are rewarded for teaching the same subjects. The determinants of rank are, unfortunately, something we were not able to explore with our data set, but would be a useful direction for future research, using productivity data such as those created by Davies, Kocher, and Sutter (2008). Are women and men equally productive or are they productive in different ways? Are they recognized equally for the work that they do?

The cross-sectional analysis demonstrates that, on average, the female and male salary structures are more similar than they once were. Yet single point estimates do not tell us the spread of a distribution; for this, we need to estimate the distribution of earnings, as was done in the third section of this paper. The distributional analysis, like the cross-sectional analysis, shows a convergence in the male and female salary structures, particularly between 1970 and 1980. It also reveals where the problems still are. In 1970, women were sharply overrepresented at the bottom of the salary distribution. By 2000, the patterns were less extreme, but women still were overrepresented at the bottom and underrepresented at the top of the distribution.

At the same time, as our distribution analysis shows, the variance of academic salaries is spreading. As we know from the cross-sectional analysis, universities with medical schools pay academics at the rank of professor relatively high salaries, and some of the discipline-specific penalties and premia are growing. We hypothesize that the university sector is becoming more fragmented both across disciplines within universities and also across universities. Exploring the determinants of the salary profiles we have documented in this paper will take more research.

Appendix

TABLE A1
Sample means by year and gender; percentage female by year

	1970			1980			1990			2000		
	F	m	% f	f	m	% f	f	m	% f	f	m	%f
Rank												
Full professor	6.5	25.1	3.4	11.0	34.9	4.8	14.9	43.3	7.6	22.2	47.6	15.3
Associate professor	21.2	32.8	8.0	36.4	40.9	12.4	35.3	35.3	19.3	39.0	32.6	31.7
Assistant professor	44.7	33.2	15.4	35.9	18.9	23.2	36.1	17.1	33.5	29.7	16.5	41.2
All others	27.6	8.8	29.6	16.8	5.3	33.3	13.7	4.2	43.7	9.1	3.3	51.8
Highest degree												
PhD	31.8	58.4	6.8	45.9	67.3	9.8	57.1	73.4	15.7	70.6	80.5	25.4
Professional	5.4	6.9	9.6	3.5	7.1	7.3	5.3	6.8	15.7	4.8	6.3	22.9
Graduate	46.2	27.5	18.5	37.0	20.1	22.7	29.0	14.9	31.7	19.5	10.3	42.4
Undergraduate	13.8	6.3	22.7	10.5	4.0	29.7	5.6	2.8	32.4	3.8	2.0	42.5
Other	2.8	0.8	32.4	3.0	1.5	23.6	3.0	2.2	25.2	1.3	0.9	35.8
First degree from												
Canada	51.8	49.4	12.4	61.6	55.6	15.0	66.0	57.2	21.6	70.7	61.2	31.0
US	14.3	14.5	11.7	18.3	17.2	14.5	14.6	14.5	19.4	11.3	12.5	26.1
UK	6.4	11.9	6.7	7.5	13.2	8.3	5.7	11.4	10.8	3.6	8.0	15.0
France	4.4	1.8	24.9	2.8	1.7	21.0	2.6	1.6	27.2	2.0	1.8	30.4
Other	23.2	22.3	12.3	9.7	12.4	11.1	11.2	15.4	14.8	12.3	16.5	22.5
Highest degree												
Canada	40.7	36.4	13.1	52.1	43.9	15.9	59.7	47.9	22.9	66.6	56.7	31.4
US	29.4	29.2	11.9	31.9	32.2	13.6	23.6	27.9	16.8	19.1	23.6	23.9
UK	5.9	13.2	5.7	6.2	13.6	6.8	5.6	12.0	10.0	5.0	9.1	17.7
France	4.9	3.4	16.2	3.9	3.2	16.6	3.6	3.4	20.2	3.1	3.2	27.3
Other	19.2	17.7	12.7	5.8	7.0	11.7	7.5	8.7	17.0	6.2	7.4	24.5

Continued

TABLE A1 *Continued*

	1970			1980			1990			2000		
	F	m	% f	f	m	% f	f	m	% f	f	m	%f
Subject taught												
Education	14.7	9.0	18.0	15.0	8.6	21.8	12.4	7.4	28.6	11.8	6.2	42.8
Fine arts	4.4	3.5	14.5	5.9	3.7	20.1	5.8	3.8	26.7	5.3	3.6	36.4
Humanities	28.3	20.5	15.7	22.6	17.0	17.5	21.1	15.4	24.7	19.0	13.4	35.5
Business/econ	2.0	6.2	4.1	2.8	8.1	5.2	5.9	9.5	12.9	6.7	10.2	20.3
Ag/bio science	8.9	7.0	14.6	7.4	7.3	14.0	6.4	7.5	17.0	6.5	8.4	23.1
Social science	13.4	13.7	11.6	18.4	15.9	15.6	19.1	15.7	22.5	20.7	15.8	33.7
Eng/applied sci	0.5	10.0	0.7	0.5	8.7	1.0	1.3	9.3	3.3	2.5	10.4	8.7
Nursing*	11.1	–	–	11.4	–	–	8.2	–	–	6.8	–	–
Health	10.7	12.7	10.2	11.7	14.6	11.4	14.0	15.4	17.9	13.8	15.3	26.0
Math/science	4.4	15.4	3.7	3.4	14.9	3.5	4.6	15.3	6.7	5.7	15.8	12.2
Other subject	1.7	2.0	–	0.8	1.3	–	1.1	0.7	–	1.3	0.9	–
Medical	63.2	61.7	12.2	58.6	58.2	13.8	54.1	56.2	18.7	52.4	55.5	26.8
Comprehensive	20.1	22.8	10.6	25.3	27.5	12.8	27.3	28.0	19.0	29.3	29.2	28.1
Teaching	9.9	8.3	13.9	12.8	10.3	16.5	14.0	11.0	23.4	15.2	12.1	32.8
Other	6.7	7.3	11.1	3.3	4.0	11.8	4.5	4.9	18.0	3.2	3.1	28.1

Year	1970		1980		1990		2000	
Gender	female	male	female	male	female	male	female	male
N	2,389	17,700	4,009	25,149	6,777	28,338	9,191	23,624
Earnings	62,902	79,687	69,093	83,026	69,869	84,301	72,357	83,992
Age	42.3	40.6	43.4	44.3	44.4	47.9	47.0	49.9

NOTES: Individuals 30–65. *For males, 'Nursing' is shown as part of 'other subject,' owing to small sample size.

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