Earnings and Employment Probabilities of Men by Education and Birth Cohort, 1982-96: Evidence for the United States, Canada and Australia

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ABSTRACT

In this paper, we analyse the earnings and employment probabilities of men by education level, birth cohort and age in the United States, Canada and Australia using a series of cross-sectional surveys for each country spanning the years 1982 through 1996. For all three countries, more recent birth cohorts of less-skilled men have experienced worse labour market outcomes than men from the same skill group but of earlier birth cohorts, ceteris paribus. In the United States, the deteriorating labour market outcomes appear as lower earnings but not lower employment probabilities. In Canada and Australia, the less skilled men from more recent birth cohorts experience lower employment probabilities and lower earnings, with the magnitude of the earnings decline by cohort being smaller than was the case for the U.S. This is consistent with the hypothesis that labour market institutions in Australia and Canada have prevented wage levels from declining sufficiently to avoid the need for reductions in employment probabilities. In the United States, wage flexibility may have removed the need for reductions in employment probabilities.

1. Introduction

In this paper, our objective is to examine differences in the labour market outcomes of recent birth cohorts of men compared with those from earlier birth cohorts by education group using data from the period 1982-1996 for the United States, Canada and Australia. We have two main questions of interest. First, has there been a decline in the labour market outcomes of more recent birth cohorts of men in each of the three countries and if so, has the magnitude of the decline varied across education groups and across countries? Second, if a decline occurred is it manifested in terms of wages in the relatively flexible labour markets of the U.S. and manifested more as changes in employment probabilities in the relatively less flexible labour markets in Australia and Canada?

The analysis uses the approach employed by Beaudry and Green (2000) in that we estimate flexible reduced form specifications to examine changes in age-earnings profiles across birth cohorts of Canadian men. They find that more recent birth cohorts of men in Canada have had lower earnings than did men from earlier birth cohorts when compared at the same age. The magnitude of these birth cohort differences was largest for high school and post-secondary men and smaller for men with university degrees. We extend their analysis to consider the age-earnings profiles of Australian and American men over the same sample period, and we use a similar approach to examine the age-employment probability profiles of men from each country.

Our paper is in a similar vein to that of Card, Kramarz and Lemieux (1999) who examine changes in the relative structure of earnings and employment of men in Canada, the United States and France over the decade of the 1980s. For each country, they employ two cross-sectional surveys - one from the beginning of the 1980s and one from the end of the 1980s – to evaluate the impact of demand shocks on wages and employment outcomes of young men across the three countries. They find that wages adjusted more to changes in demand conditions in the U.S. than in Canada and France, and this is consistent with their hypothesis that the U.S. labor market is more flexible than the Canadian or the French labor markets. In contrast to this hypothesis, they find that employment was no more responsive to the change in demand conditions in Canada and France than in the United States.

Our main objective is to characterize the labour market performance of recent birth cohorts of men by skill group in three similar industrialized countries. Because of institutional differences across the labour markets in the three countries, we evaluate both earnings and employment probabilities for men from different educational backgrounds and birth cohorts. This allows us to investigate whether changes in the overall labour market conditions lead to different labour market outcome responses across the three countries. Given this objective, we adopt a reduced form approach to estimation that facilitates comparisons of general patterns in earnings and employment probabilities across countries. Our approach does not allow us to test that specific sources of hypothesized labour market inflexibility in Canada and Australia have led demand changes to be manifested more in employment probability adjustment and less as wage adjustment. Nonetheless this reduced form approach yields important insights into this issue.

Our results indicate that the labour market outcomes of more recent birth cohorts of less skilled men have deteriorated in all three countries. For high school and post-secondary educated men in all three countries we find evidence of lower earnings at the same age than were received by men with the same education level from earlier birth cohorts. The earnings cohort effects are most pronounced in the U.S. case. In addition, we find that employment probabilities display a similar pattern of deteriorating outcomes for more recent birth cohorts of men with education below the university level in Canada and Australia but not in the United States, where no clear cohort pattern is observed. This evidence is consistent with the hypothesis that the U.S. labour market is more flexible than the Canadian and Australian labour markets. Wages appear to adjust down, avoiding the need for decreases in employment probabilities for more recent birth cohorts of less-skilled men when labour market conditions deteriorate.

2. Institutional Differences Across the Three Countries

A number of authors have written on differences in the labour market institutions between Canada and the United States.[1] Rather than review this literature, we will highlight key differences that are relevant to the analysis of this paper. Benefits available to the unemployed are more generous in Canada than in the United States, allowing displaced workers to search for work for longer. A relatively generous unemployment compensation system also exists in Australia. Key differences are that the Australian unemployment benefits are not increasing in the person's past income as they are (to a maximum) in Canada. However, unlike in Canada, there is not a termination date for receipt of benefits while unemployed in Australia.

In addition, both Canada and Australia have higher rates of union contract coverage than is the case in the United States. Card, Kramarz and Lemieux (1999) state that union membership in Canada was 50 percent higher than in the United States in 1990 and over 100 percent higher than in the United States in 1995. In Australia, virtually all wage and salary workers are covered by a union-negotiated employment contract, even though union density had fallen to 35% by 1994.

A third possible area of difference in labour market policy is in terms of minimum wage rates. Baker, Benjamin and Stanger (1999) present graphs of the minimum wage across nine of the provinces of Canada from 1975-1993. The general pattern is of declining

[1] See for example Card and Freeman (1993) and Card, Kramarz and Lemieux (1999). See also Baker, Benjamin and Stanger (1999) for a recent study of the impact of minimum wages on employment of young workers in Canada.

minimum wages over most of the period with a small increase in the minimum wages in Alberta, British Columbia, Ontario and Quebec near the end of the period. Card, Kramarz and Lemieux (1999) present a graph of the minimum wage relative to the average manufacturing wage in the United States and Canada over the period 1966 to 1994. Over our sample period, these two series move fairly closely together.

The base wages of most workers in Australia have until recently been determined within a centralized system of national, state or industry Awards.[2] Thus, there is no overall minimum wage rate but rather a set of industry and occupation specific minimum wages. For comparison purposes, we have derived a wage series that is based on the lowest award wage levels in the Australian economy - the Award wage rates applicable to unskilled labourers in the Textile, Clothing and Footwear industry. This series is then deflated by the average ordinary time earnings of men in manufacturing to yield a rough measure of a relative minimum wage rate similar to the one generated for Canada and the United States in Figure 2 of Card, Karamarz and Lemieux (1999).[3] We plot our Australian relative minimum wage series in Figure 1. Over the 1982-96 period, relative minimum wage rates in Australia generally declined. From Figure 1 of Card, Kramarz and Lemieux (1999), relative minimum wage rates in the United States declined from 1982 until 1989 and then rose after that. The Canadian relative minimum wage was fairly constant over the period, rising only in the early 1990s. The Australian relative minimum wage rate is considerably higher than in the United States and Canada over the entire 1982-1994 period. Specifically, the Australian minimum

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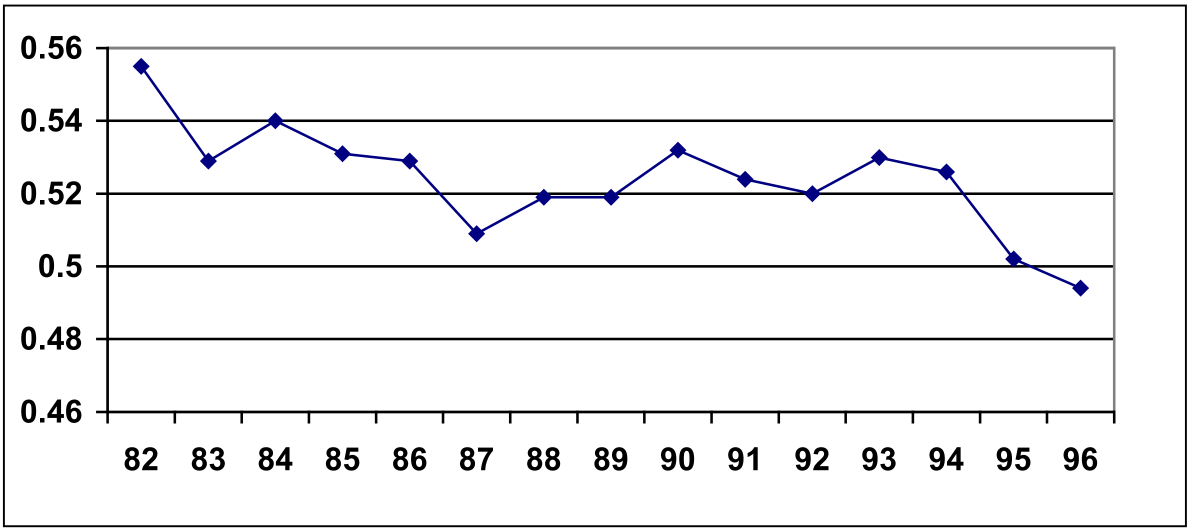
[2] Under the Australian Award system, minimum terms and conditions of employment are specified for most job classifications, in effect providing a series of minimum wage rates. While the centralized wage determination system provided by the Prices and Incomes Accord was abolished in 1996 and replaced by a system of enterprise bargaining, the Awards continue to provide minimum wages and conditions that enterprise bargains must meet.

[3] See the appendix for details of the construction of the minimum wage series for Australia.

wage rate varies between 55 percent and 49 percent while the United States relative minimum wage rate series varies from approximately 39 percent in 1982 to approximately 37 percent in 1993. Similarly, the Canadian relative minimum wage rate varies from approximately 35 percent in 1982 to approximately 37 percent in 1993.[4]

Figure 1

Derived Minimum Wage Relative to Average Wage of Men in Australia: 1982-1996



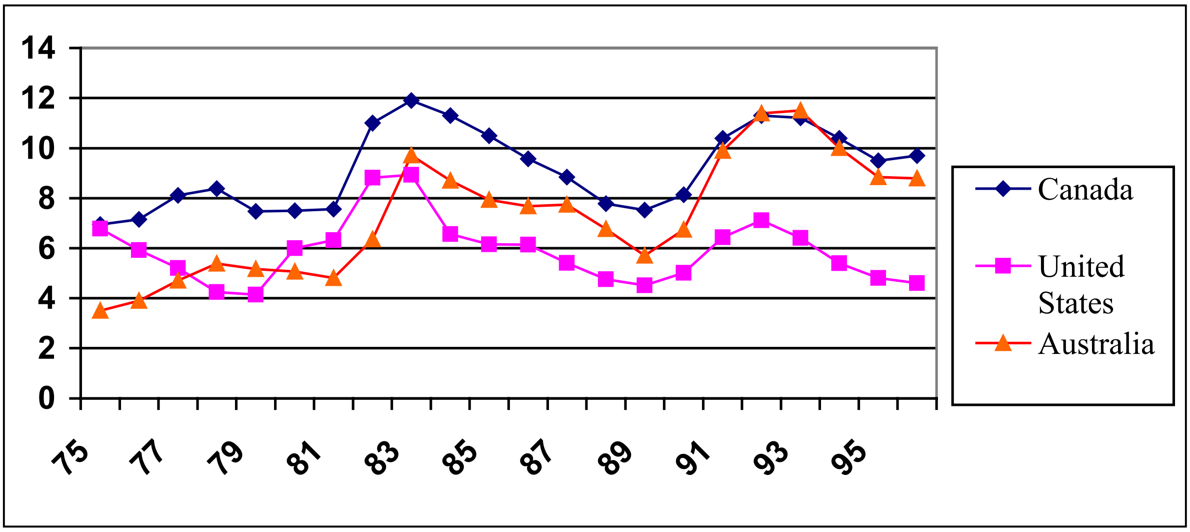
Source: Authors' calculations. See Appendix 1.

In Figure 2, aggregate male unemployment rates are presented for each of the three countries over the period 1975 to 1996. A similar business cycle pattern is apparent with recessions occurring in the early 1980s and early 1990s in each country. Taking the late 1970s as a starting point, the Australian and American unemployment rates were similar in magnitude at roughly five percent and the Canadian unemployment rate was higher at around

[4] A related source of difference might be the degree of centralization in wage determination. Over our sample period, Australian wage determination was relatively centralized. However, there is some contention in the literature about the links between centralized wage determination and the responsiveness of wages to

eight percent. Over time the Australian unemployment rate and the Canadian unemployment rate have converged, leaving a difference of four or five percentage points when compared with the U.S. unemployment rate by 1996. This preliminary evidence supports the idea that a deterioration in labour market outcomes may be manifested in terms of drops in employment probabilities in Canada and Australia rather than in wage declines. Consequently, wage rates may be less responsive to changing economic conditions in Canada and Australia than in the more deregulated labour markets of the United States.

Figure 2
Aggregate Male Unemployment Rates
in the United States, Canada and Australia: 1975-1996



Source:
1. The Australian unemployment rates are taken from the Australian Bureau of Statistics Labour Force, Australia 6204.0.
2. The United States unemployment rates are taken from Bureau of Labor Statistics `Civilian Labor Force 20 years and over male' LFS21001701.
3. The Canadian unemployment rates are taken from Statistics Canada, *Canada, Labour Force Characteristics*, D980404.

demand shocks. For example, Coelli et.al. (1994) find Australian wages are strongly correlated with factors influencing the demand for labour.

The balance of the paper proceeds as follows. In Section 3, the econometric specifications are presented. In Section 4, the data are described and the samples used in estimation are defined. Results from the earnings equation estimation for each of the three countries are discussed in Section 5. The results from the analysis of employment probabilities for the three countries are discussed in Section 6. In Section 7, results are discussed from a sensitivity analysis involving estimation over the period 1982-1989 as opposed to the full sample period. Concluding remarks are discussed in Section 8.

3. Econometric Specifications

The analysis involves least squares estimation of log weekly earnings equations and probit estimation of binary outcomes models over employment versus non-employment (either unemployment or non-participation in the labour force). In each case the same specification of the explanatory variables is used:

$$Y_i = \beta_{c1} C_i + \beta_{c2} C_i^{\,2} + \beta_{ca} C_i \times A_i + \beta_{a1} A_i + \beta_{a2} A_i^{\,2} + \beta_{a3} A_i^{\,3} + \beta_u U_t + \varepsilon_i \qquad (1)$$

where Y_i is the natural logarithm of the current weekly earnings of individual *i* for the case of the earnings equation estimation; alternatively, Y_i is a latent index for the case of the probit estimation of employment incidence.[5] In the latter case, the index is positive if the individual is employed and negative if the individual is either unemployed or not in the labour force.

The equation employs the specification used by Beaudry and Green (2000). A continuous cohort variable, C_i, is defined giving the year of birth of the individual (ie. a

[5] We explore the sensitivity of the results to the inclusion of demographic characteristics. These results are described below. We also estimate more flexible specifications of the age cohort specifications to explore the sensitivity of the results. In particular, we allow for separate cubic age profiles by birth cohort for each country/education sub-sample. The main results of the paper are not sensitive to the use of this more general specification.

person born in 1955 would have a value of the cohort variable of 55). The cohort variable appears in linear form, quadratic form and as an interaction with the continuous age variable, A_i. The age variable appears in linear, quadratic and cubic forms. Finally, the current aggregate unemployment rate, U_t, appears as a control for period effects.

The earnings and employment probability models are estimated separately for each country by education group. The sample is divided by country into men with: 1) high school education or less, 2) some post-secondary education but no university degree and 3) a university degree.

The identification of birth cohort, ageing and period effects is achieved under the restrictions that period effects are the same for everyone and can be captured parametrically through the current unemployment rate. Given this restriction on the form of period effects, the use of multiple cross-sections for each country allows for the separate identification of the cohort effects from the ageing effects. The interaction variable between the cohort variable and the age variable allows for the cohort differences to vary with age so that cohort differences need not be fixed differences lasting throughout the workers' careers.

However, it is possible that period effects may vary by age group and education group.[6] In particular, it is likely that period effects may have a larger impact on more recent labour market entrants. We do not factor that into the empirical model; however, we do allow for this interpretation when discussing the results. It should be stressed that since our specification restricts period effects to be the same for all workers within the country/education group sub-sample, any differences in the impact of the period effects by group will be absorbed into the coefficient estimates on the birth cohort and/or age

[6] We have found this to be important in our previous work using the SCF data (see McDonald and Worswick 1998, 1999).

variables.[7] In Section 7, we investigate the sensitivity of the estimated models to the survey years used.

4. The Data and Estimation Sample

The data used in the estimation come from three sources: 1) the Australian *Income Distribution Surveys* (IDS) of the Australian Bureau of Statistics for the years 1982, 1986, 1990, 1994, 1995 and 1996;[8] 2) the *Survey of Consumer Finances* (SCF) of Statistics Canada for the years 1982, 1983, 1985 through 1996, and 3) the March Supplements of the *Current Population Survey* (CPS) of the U.S. Bureau of the Census for the years 1982 through 1996. A set of sample weights is provided in each data set and is used in the estimation to enable generalizations of the results to the relevant population. These three surveys ask similar questions regarding economic and demographic information. Each is a supplement to the monthly national labour force survey.

For each country, the sample is restricted to men between the ages of 25 and 55 in each of the surveys. We choose to focus on men since the determinants of employment versus non-employment are more likely to be related to factors such as presence of young children for women than for men. An analysis of employment probabilities of women would involve addressing the large increases in labour force participation for more recent birth cohorts of women for all three countries. Analyzing only men allows us to focus more clearly on the role of labour market institutions in the discussion and to reduce the number of

[7] For example, consider the case of a recent birth cohort entering the labour market during a severe recession and this recession effect being an abnormally large negative impact on their earnings and employment probabilities relative to earlier birth cohorts. This might result in a large birth cohort effect for this group relative to the earlier cohorts implying lower earnings and higher employment probabilities at the same age ceteris paribus.

[8] These are all of the publicly available IDS files that were available at the time of commencement of this project. For the first four surveys used, data were collected in the fourth quarter of the Calendar year. For the

versions of the various models in the discussion. The age restrictions are designed to focus the analysis on men that are likely to have a strong labor force attachment. Men under the age of 25 may be involved in post-secondary studies, and therefore by excluding them, we avoid problems associated with differential sample selection by educational attainment. A similar problem relating to early retirement decisions is mitigated by focussing on workers under the age of 56.

Immigrants are excluded from the Australian and Canadian samples in order to consistently define the underlying population analysed in each cross-sectional survey. Unfortunately, the March CPS data do not contain immigrant information prior to 1994; therefore, we are unable to exclude immigrants from the American sample.[9]

After the restrictions we are left with samples of 470 634 for the United States, 271 640 for Canada and 23 607 for Australia. The difference in sample size between the American data and the Canadian data is due to the larger sample surveyed in each year in the American data and the fact that the Canadian survey was not carried out for the 1984 year. The Australian sample size is less than ten percent of the Canadian data due to the much smaller sample size and the infrequency of surveys over the 1982-1996 period.

In the analysis of employment probabilities, all men in the age range are studied. In the analysis of weekly earnings, only men with positive earnings from wages and salaries who were not self-employed are included in the sample used in estimation.[10]

1995 and 1996 surveys, households were surveyed continuously over the financial year. For example, for the 1996 survey, the financial year commenced in July 1996 and ended in June 1997.

[9] As a sensitivity check, we re-estimated the various models for Australia and Canada with the immigrants in each case included in the sample. While there were small changes in the magnitude of the estimates, the general pattern of cohort effects in each case were unaffected.

[10] Estimation was also carried out using hourly wages (imputed from weekly earnings and hours worked per week). The results were qualitatively unchanged. We have adopted the weekly earnings measure since the hours of work information is top-coded at 35 and 65 in the Australian and Canadian data, respectively.

In Table 1, sample means of key variables are presented. Mean weekly earnings are presented in the local currency and denote a positive return to increased education in each country. Mean weekly earnings for university graduates are 68.6, 42.4, and 54.8 percent higher than mean weekly earnings of men with high school education or less in the United States, Canada and Australia, respectively. Similarly, mean weekly earnings of men with post-secondary education but no university degree are 20.9, 10.3 and 16.2 percent higher than mean weekly earnings of men with high school education or less in the United States, Canada and Australia, respectively.

Employment incidence does not differ greatly on average across the three countries when the education level is held the same. In particular, at the university level the employment probabilities vary from a low of 92.7 percent for the United States to a high of 95.4 percent for Australia. Larger differences are present at the other education levels with Canada having the lowest employment probabilities followed by the United States and then Australia for both high school men and men with post-secondary education below the university degree level.

One important difference is in the fraction of men in each education level in each country. The fraction of men with university degrees is 26.8 percent in the United States, 13.2 percent in Canada and 12.1 percent in Australia. The fraction of men with post-secondary education below the university degree level is similar in the United States and Canada at 21.6 and 23.1 percent, respectively; however, it is almost twice that size at 42.4 percent in Australia. While there are strong similarities in the education systems and programs across the three countries, these differences in the sample proportions may indicate different selection processes into the different education programs. This may be important in explaining differences in labour market outcomes across the three countries for men with a particular level of education.

Table 1

Sample Means by Education Level:
The United States, Canada and Australia

	Country	High School	Post-Secondary	University Degree
Weekly Earnings	United States	549.75	664.44	927.08
	Canada	717.16	791.12	1021.45
	Australia	635.46	738.48	983.62
Employment Incidence	United States	0.8135	0.8751	0.9270
	Canada	0.7838	0.8667	0.9310
	Australia	0.8385	0.9233	0.9544
Age	United States	37.32	36.61	37.76
	Canada	37.95	36.08	37.42
	Australia	38.09	38.06	37.24
Sample Size (fraction of national sample in parentheses)	United States	243 025 (0.5164)	101 724 (0.2161)	125 885 (0.2675)
	Canada	173 079 (0.6372)	62 634 (0.2306)	35 927 (0.1323)
	Australia	10 766 (0.4561)	9 997 (0.4235)	2 844 (0.1205)

Note:
1. The earnings figures are in the 1995 local currency.
2. The labour force status information in the public use March CPS file was not available; therefore, we excluded the 1994 survey when carrying out the employment probability estimation. In this case, the sample size drops to 229 653, 93 956 and 117 641 for high school or less, post-secondary but no university degree and university degree men, respectively.

5. Male Earnings by Education, Birth Cohort and Country

5.1 The United States

The regression output from the majority of the models estimated in the paper is presented in Table 2. Because of the large number of parameter estimates generated in the various regressions generated in the analysis of this paper, we focus the discussion on the predicted age-earnings profiles by cohort for each country. It should be borne in mind that meaningful comparisons of cohort effect magnitudes across countries implicitly require that the nature of demand shocks over the sample period be comparable across the three countries. Card, Kramarz and Lemieux (2000) argue that this has been the case through the 1980s for Canada, the United States and France, and Australia has also had a similar experience. Nonetheless, this caveat should be borne in mind.

In Figure 3, age-earnings profiles of American men are presented by education level and by birth cohort. The values of age over which the curves are plotted correspond to the ages observed in the sample period 1982-1996 for the birth cohort. For men with high school education and post-secondary education, a clear pattern of differences across birth cohorts in earnings at the same age is apparent. More recent birth cohorts have had lower earnings than their predecessors at the same age and these differences appear to be permanent. The size of the differences is larger for the high school men compared with the post-secondary men. In contrast, the predictions of the third panel of Figure 3 indicate that differences in average earnings by birth cohort at the same age are not present for American men with university degrees.

In Table 3, p-values of tests of two restrictions on the birth cohort effects are presented for all of the models in the paper. The first restriction is that there are `no differences' across birth cohorts. In terms of equation (1), this restrictions sets $\beta_{c1}$, $\beta_{c2}$ and $\beta_{ca}$ all equal to zero. The second restriction allows for the possibility of birth cohort differences but restricts them to be `fixed differences' meaning that the cohort differences do not vary with the age at which the comparison is made. In terms of equation (1), this restriction sets β_{ca} equal to zero. We report p-values for the base case specifications of the model as shown in equation (1) and for a richer set of models where mainly demographic variables are included in the model. We do not present graphs from the models where demographic variables are included to economize on space. The results do not change qualitatively when the more general model is used.[11]

As can be seen in Table 3, the restriction of no birth cohort differences can be rejected for men from all three of the education groups for the United States. However, as can be seen in the figure, they only appear to be economically meaningful in magnitude for men with education below the university degree level. The significance of the restriction that cohort effects are fixed varies by education group and by whether covariates are included. From Figure 3, it does not appear that the size of the differences across birth cohorts varies greatly with age.

Overall, the results indicates that while the returns to a university degree have been rising across time, it is due to a decline in the earnings of less skilled workers across birth cohorts rather than a growth in the returns to skilled (university-educated) workers, across birth cohorts.[12] The fact that deteriorating labour market outcomes for relatively unskilled men are being manifested (at least in part) as lower wages for more recent birth cohorts is consistent with the idea that wages are flexible in the U.S. and can adjust down in response to poor labour market conditions. As is shown below, the magnitude of earnings cohort effects for less skilled workers is largest in the U.S. case when compared with the results for Canada and Australia where it is argued that institutional features may prevent downward wage adjustment.[13]

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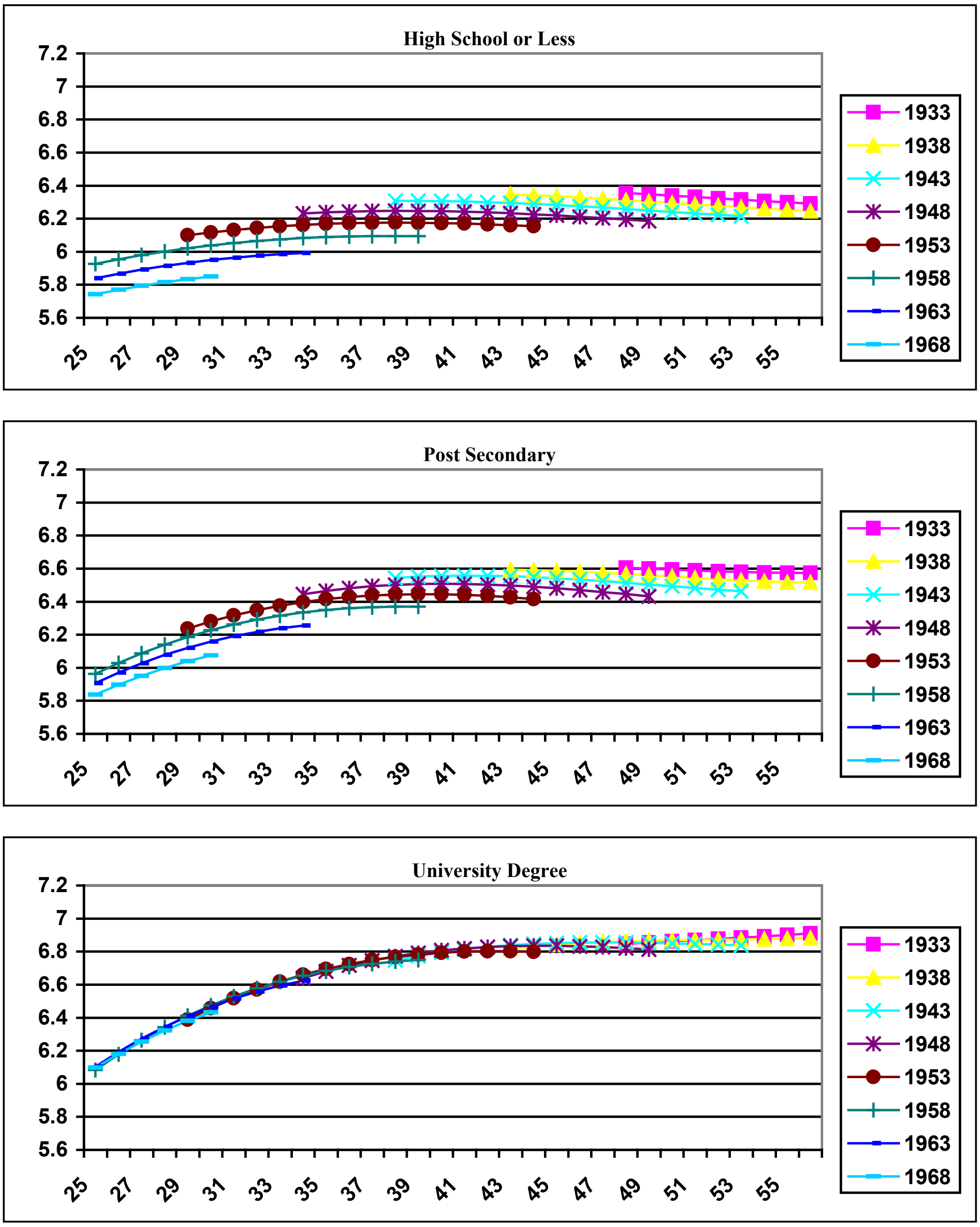
[11] Demographic controls include dummy variables for ethnicity, and not surprisingly, these variables are highly significant. Estimating the model separately by ethnic group reveals substantial differences in the patterns of earnings and employment cohort effects for blacks and hispanics compared to whites. We are investigating this issue in a companion paper and so exclude discussion of the main results here. It should be noted however that restricting the sample of American males to whites only has little impact on the results reported in this paper.

[12] See Bound and Johnson (1992), Murphy and Welch (1992), DiNardo, Fortin and Lemieux (1996), Katz and Murphy (1992), Juhn, Murphy and Pierce (1993) and Borjas and Ramey (1994) for key articles in this area.

[13] The earnings analysis for each country uses real weekly earnings in the local currency in 1995. The dependent variable is the natural logarithm of real weekly earnings. The levels of earnings are not directly

Figure 3

Age-Earnings Profiles of in the United States by Education and Birth Cohort



Note:

1. The dependent variable is the natural logarithm of weekly earnings.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

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comparable across the countries; however, the dispersion in the cohort effects is comparable since they can

5.2 Canada

Figure 4 presents three diagrams that are equivalent to those of Figure 3 but are generated using the SCF data for Canadian men. Given the institutional differences between Canada than the United States outlined above, we would expect that deteriorating labour market outcomes for more recent birth cohorts of Canadian men to be manifested in both lower weekly earnings and lower employment probabilities. The more generous unemployment benefits available in Canada than the United States may make unemployment less costly for men in Canada and lead to more prolonged job search. The institutional differences are more likely to have a larger impact near the bottom of the skill distribution; therefore, we expect to see larger differences in behaviour between men with high school level education or less in Canada compared with similarly educated men in the United States.

For men with only high school education, the earnings of more recent birth cohorts have been lower than those of earlier birth cohorts over the early part of their careers. For men with post-secondary education and men with university degrees, as was found for the men with high school education or less, more recent birth cohorts with post-secondary education have lower earnings than did earlier cohorts at the same age. The differences are larger in magnitude for the post-secondary men compared with the high school men or university men and do not vary with age indicating that these are fixed or permanent differences across birth cohorts.[14]

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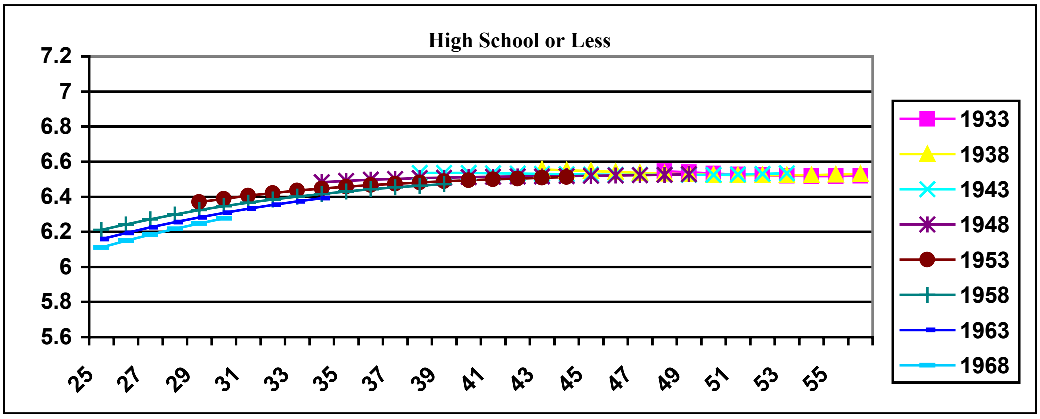
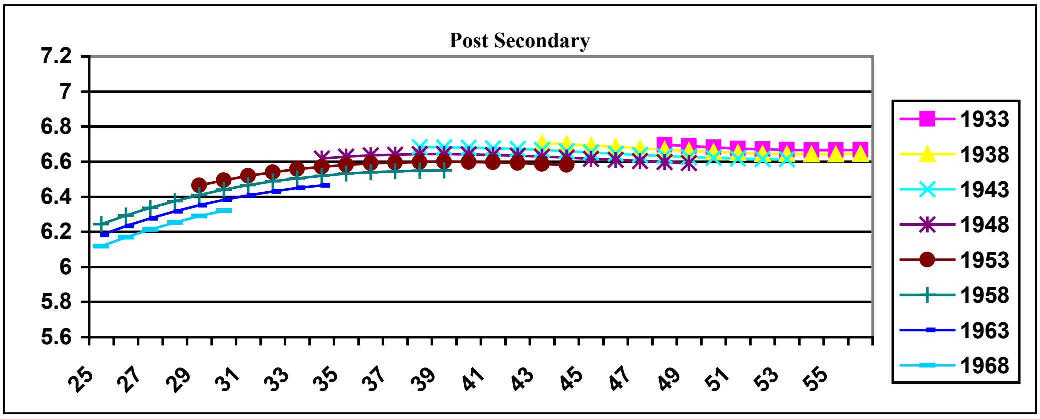
be interpreted as percentage differences from the mean.

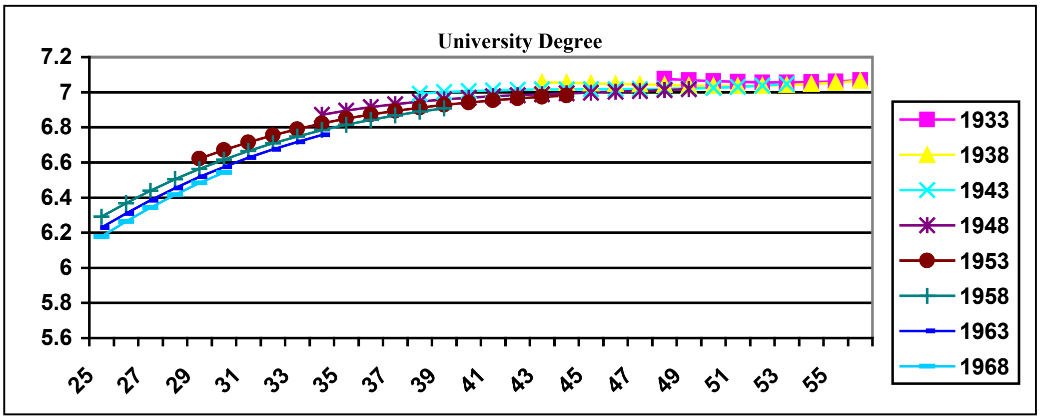
[14] The results for all three education groups of Canadian men are broadly consistent with what was found by Beaudry and Green (2000) using the same model and SCF data over the 1971 through 1993 period.

Figure 4

Age-Earnings Profiles of Men in Canada by Education and Birth Cohort







Note:

1. The dependent variable is the natural logarithm of weekly earnings.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

In Table 3, the cohort differences are significant for all three of the education groups whether demographic controls are included or not. The restriction of fixed differences can be rejected at the high school or less level but the test result is sensitive to the inclusion of demographic variables for the models estimated over the sample of men with post-secondary and university level education.

It is worth noting that the existence of cohort effects for university graduates in Canada is in contrast with what was found in Figure 3 for the U.S. where no cohort effects were present.[15] Also, the size of the cohort differences found for high school workers in Canada are much smaller than what was found for the United States. If the same forces are driving the deterioration in labour market outcomes of less skilled workers in Canada and the United States, then the results would indicate that the effects are being manifested in terms of larger falls in earnings in the United States than in Canada across birth cohorts.

5.3 Australia

Given the centralized wage setting system that existed in Australia over the 1980s and early 1990s, we might expect less wage variability across birth cohorts then what we found in the American and Canadian data. Also, given that the Australian unemployment benefit system is more generous than what exists in the United States, it may be expected that wages do not drop by as much in Australia in response to a fall in demand for less-skilled labour as would be the case under the same fall in demand for less-skilled workers in the United States.

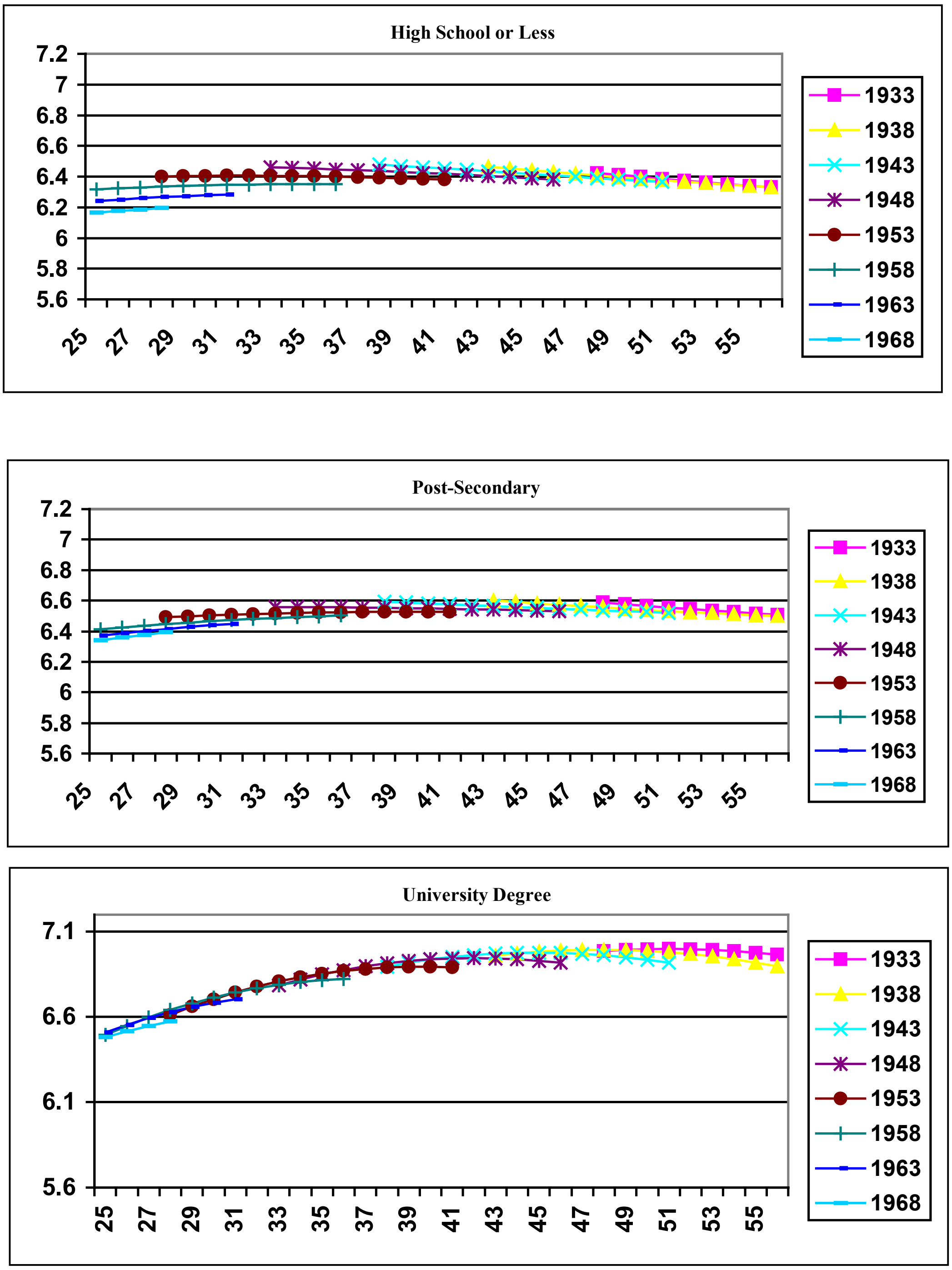
[15] Murphy, Riddell and Romer (1998) argue that the college-high school wage premium did not increase as much in Canada as in the United states during the 1980s and 1990s due to the large increase in the relative supply of university graduates in Canada over the period. This may explain the existence of cohort effects over the early part of the career for more recent birth cohorts with university degrees in Canada.

Figure 5 contains results from the estimation of the earnings equations using the Australian data. Interestingly, the institutional wage setting has not preevented the existence of cohort effects in weekly earnings for Australian men with high school or less education and to a lesser extent for Australian men with post-secondary education but no university degree. This may reflect the Awards system's ability to repond to market forces in determining occupation-specific Award minimum wages by skill group.[16] The results are similar to what was found for the U.S. in that the weekly earnings cohort effects are particularly pronounced for the less educated workers.[17] However, it is important to note that the magnitude of the cohort differences in weekly earnings for Australian men with education below the university degree level are smaller than what are found in Figure 3 for American men.

[16] This could occur through over-Award wage payments negotiated at the industry or enterprise level. This so-called two-tier bargaining system commenced in 1987 following a period of partial wage indexation. It is also possible that this was facilitated by the move to enterprise bargaining coupled with the real wage restraint near the end of the Accord period. See Dabscheck (1995) for further discussion.

[17]The p-values of Table 3 indicate that the birth cohort effects are significant for the high school men and the post-secondary men whether demographic variables are included or not. The cohort effects are not significant for university men with or without the demographic variables. The fixed birth cohort effects restriction can generally not be rejected with the exception of the case of post-secondary men when demographic characteristics are included where the p-value is .0455.

Figure 5
Age-Earnings Profiles of Men in Australia by Education and Birth Cohort



Note:
1. The dependent variable is the natural logarithm of weekly earnings.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

6. Employment Probabilities

The next stage of the analysis focuses on employment probabilities. Probit estimation is used to generate predicted employment probabilities by age, education and birth cohort for each of the three countries. Our interest is in investigating whether differences in labour market institutions across the three countries have led to differences in employment probabilities across different educational groups. Our prior belief is that the negative demand shocks experienced by low wage workers in Australia and Canada are likely to be manifested by lower probabilities of employment compared with low wage workers in the United States or other workers.

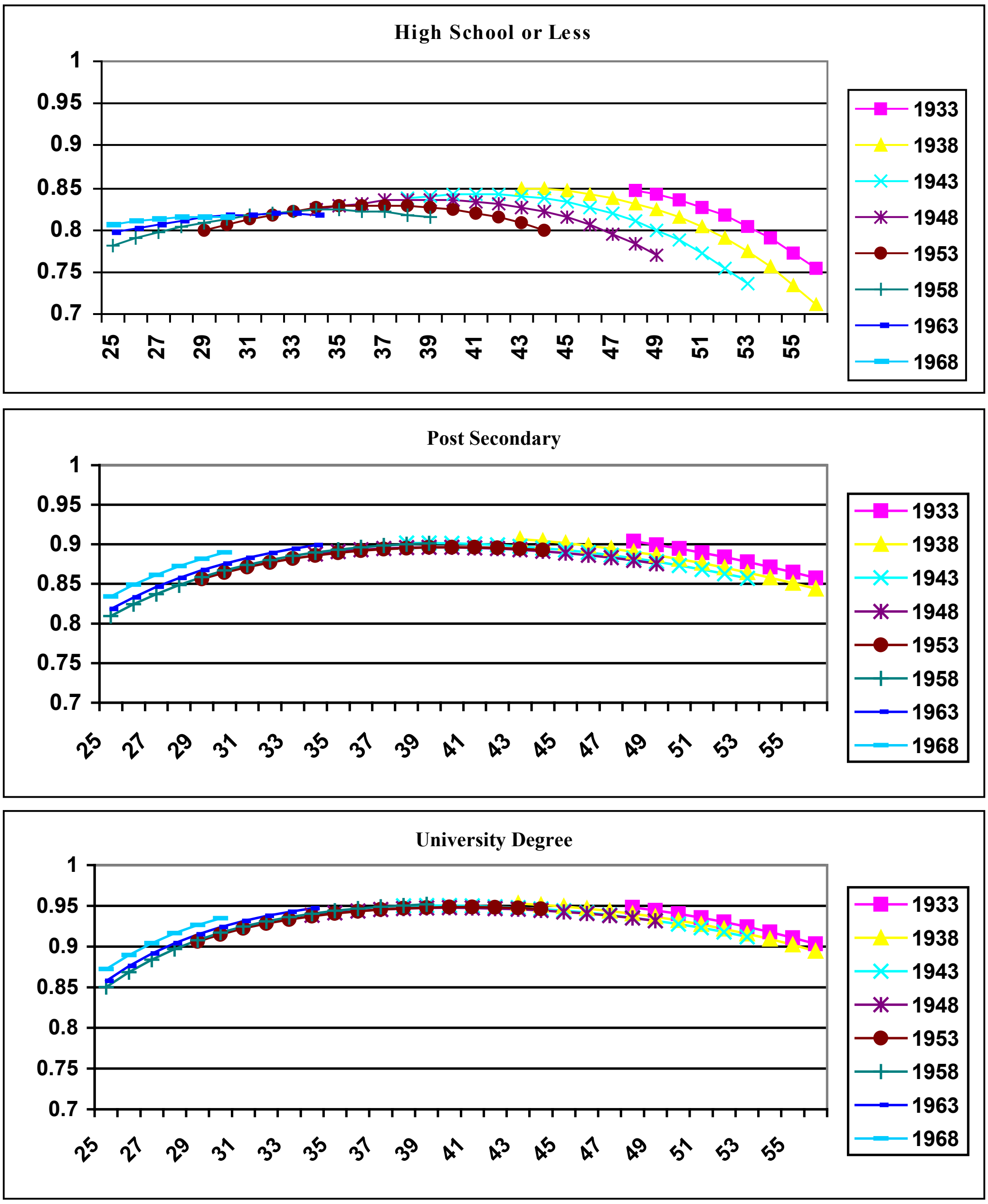
6.1 The United States

The three panels of Figure 6 contain predicted employment probabilities generated from Probit estimation using specification (1) over the samples of American men with: 1) high school education or less, 2) post-secondary education but no university degree, and 3) a university degree, respectively. From Figure 6, it is apparent that employment probabilities increase with age over the early part of the worker's career but then decrease with age near the end of the career. There appear to be differences in employment probabilities across birth cohorts at the same age; however, they do not appear to follow a simple pattern in all three of the panels. The cohorts born in 1953 and earlier display a pattern of declining employment probability at the same age across cohorts. For the more recent cohorts, the opposite pattern emerges where more recent birth cohorts have had higher employment probabilities at the same age.[18]

[18] The p-values presented in Table 3 indicate that the birth cohort effects are significant in all three cases whether demographic controls are included or not.

Figure 6

Age-Employment Probability Profiles of Men in the United States
by Education and Birth Cohort



Note:

1. Profiles generated from probit estimates over probability of employment versus non-employment.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

Overall, the American employment probability results are not consistent with a deterioration in employment probabilities for more recent birth cohorts of men compared with men from earlier cohorts holding age and education level the same.

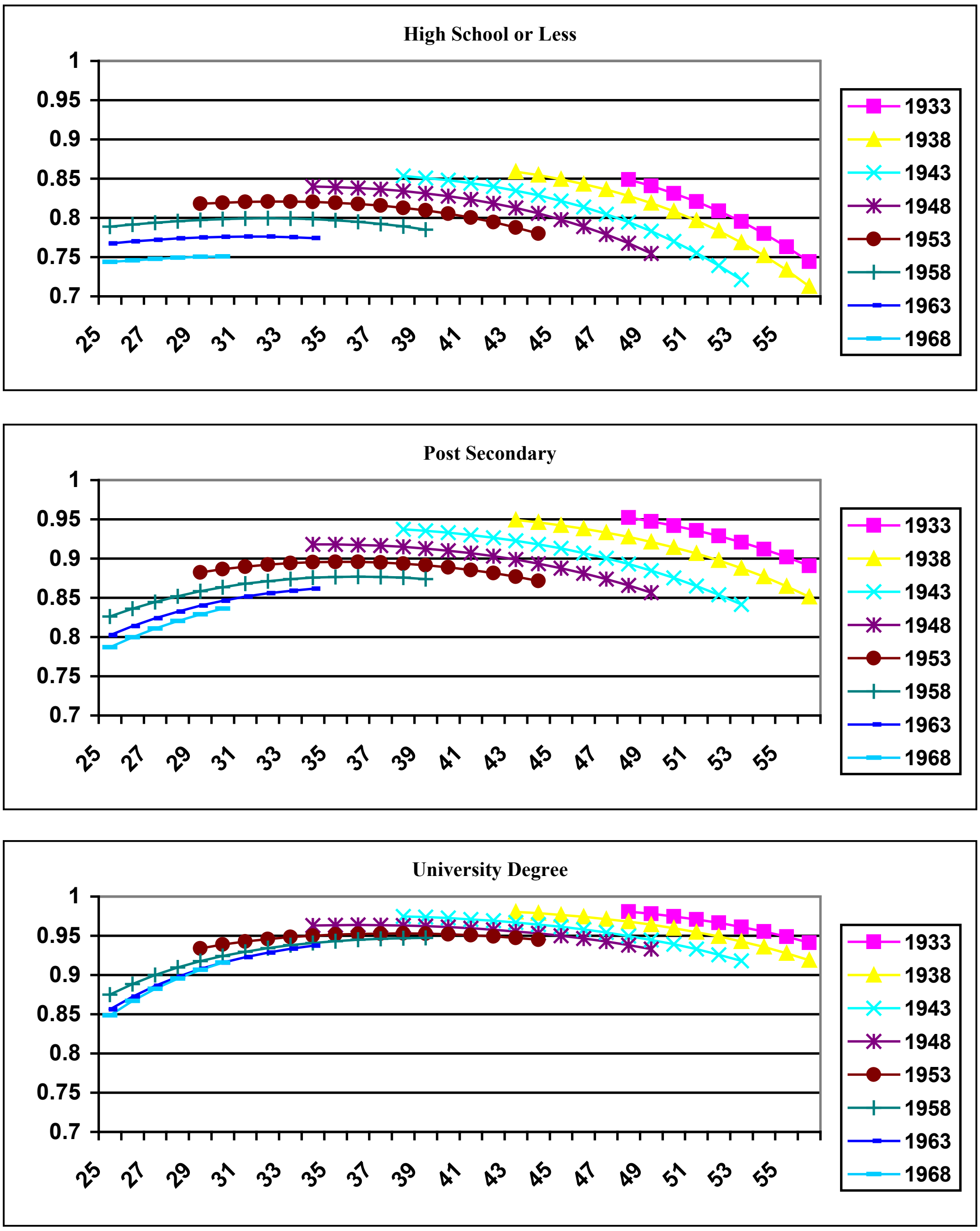
6.2 Canada

The analysis is repeated using the Canadian data, and the employment probability-age profiles by cohort are presented for the three education groups in the three panels of Figure 7. Unlike the American results, the Canadian results contain strong evidence that more recent cohorts have experienced lower employment probabilities than did earlier cohorts at the same age. This is true for men from each of the education categories.[19]

This indicates that an important difference in the type of labour market adjustment experienced by recent birth cohorts in Canada and the United States in response to the deterioration of the labour market in each country for new entrants. The pronounced employment cohort effects for Canadian men with education at or below the high school level is consistent with the idea that institutional factors such as the generous unemployment benefit system and the relatively high degree of unionization in Canada have prevented a decline in wages in response to the deterioration of the labour market, leading instead to an adjustment along the employment dimension. It is also interesting to note that the magnitude is largest for men with education at or below the high school level.

[19] The p-values of Table 3 indicate that the birth cohort effects are significant for men in each of the three educational groups whether demographic characteristics are included or not. Also, in each case we cannot reject the restriction that the cohort effects do not vary with age.

Figure 7
Age-Employment Probability Profiles of Men in Canada
by Education and Birth Cohort



Note:

1. Profiles generated from probit estimates over probability of employment versus non-employment.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

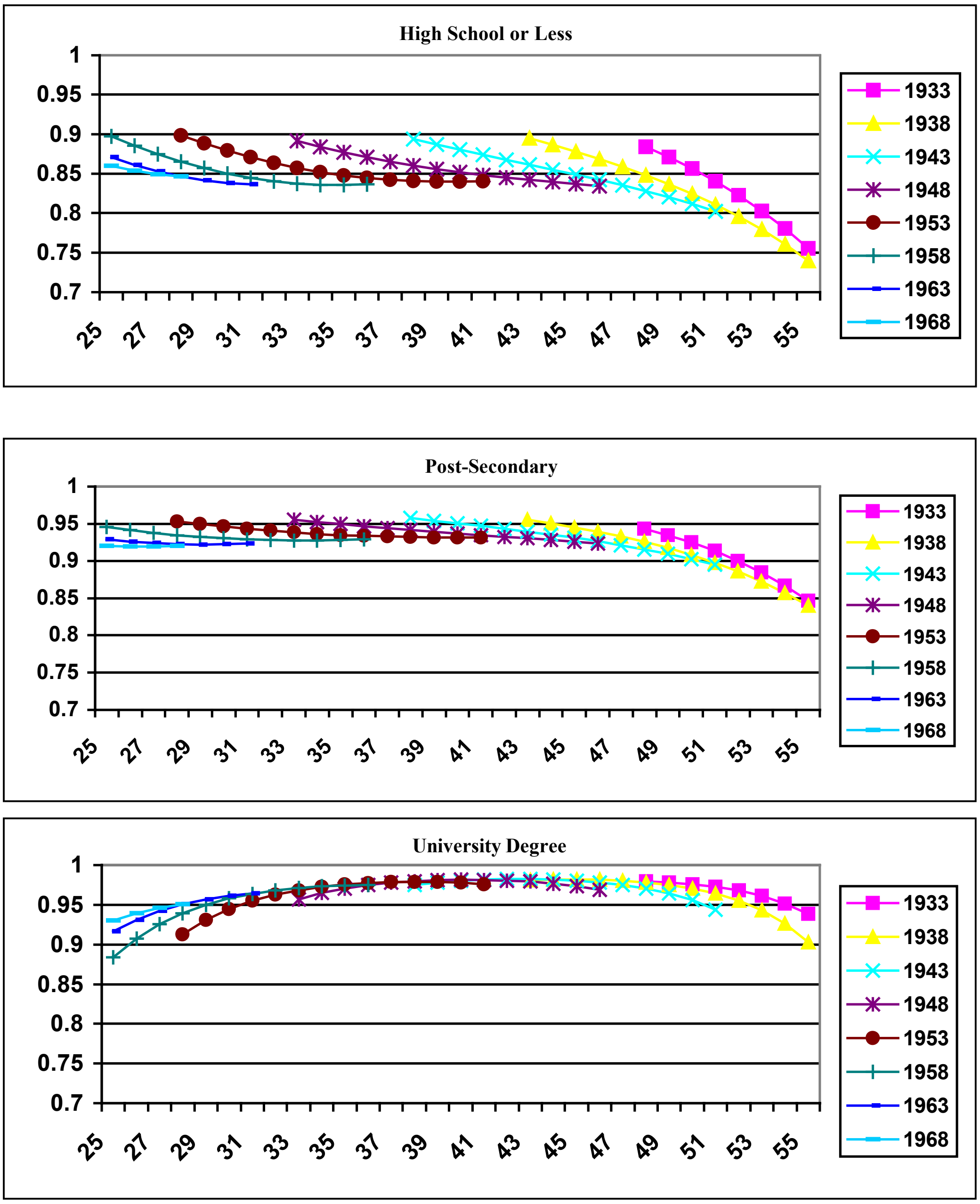
6.3 Australia

The three panels of Figure 8 present the predicted employment probability-age profiles by birth cohort and education level for Australian men. In the first two panels of Figure 8, for Australian men with education below the university degree level, employment birth cohort differences are found. Men from more recent birth cohorts have lower employment probabilities then did men from earlier birth cohorts at the same age.[20] This result is similar to what was found for these two education groups in Canada. Unlike in the Canadian case, the predicted profiles indicate that the cohort effects are not permanent. However, the restriction that the cohort effects are fixed can generally not be rejected. For Australian men with university degrees, there is a similar birth cohort pattern to what was found in Figure 6 for the United States.

The Australian employment patterns are consistent with employment being an important dimension along which poor labour market conditions are manifested for new cohorts of labour market entrants. The evidence indicates that the high minimum wage system implemented under the Awards system coupled with the relatively generous unemployment benefit system that exists in Australia may be preventing downard adjustment in wages, leading instead to adjustment in terms of employment probabilities. However, given the evidence of modest wage cohort effects for Australian men without university degrees, it appears that the adjustment to deteriorating labour market conditions for new entrants is along both the wage and employment dimension.

[20] The p-values of Table 3 indicate that the birth cohort effects are significant for men with high school or less education in either version of the model. For the post-secondary men, the cohort effects are significant only if the demographic variables are not included and the cohort effects are not significant for men with university degrees under either version of the model.

Figure 8
Age-Employment Probability Profiles of Men in Australia
by Education and Birth Cohort



Note:
1. Profiles generated from probit estimates over probability of employment versus non-employment.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

7. Sensitivity of the Results to the Choice of Sample Period

If the assumption that period effects have the same impact on all workers within an education category for a given country is invalid, then our estimates of age and cohort effects on earnings and employment probabilities will include the effect of changing period conditions such as changing macroeconomic conditions. In this section, we investigate this possibility for Canada and the United States.[21]

In Figure 9, we present the equivalent diagrams for High School men in Canada from Figure 4 and Figure 7 where the estimation is carried out over the sub-sample of Canadian men with education at or below the high school level from the 1982 through 1989 surveys. By dropping the second half of our sample years we are able to see whether the results are sensitive to removing observations from the recession years of the early 1990s and the strong growth period that followed.

The earnings regression results are presented in the first panel of Figure 9. They indicate smaller earnings cohort effects than are found in Figure 4 using the full sample. However, the general pattern is similar. In the next panel of Figure 9, we see that the pronounced employment cohort effects of Figure 7 are reversed, implying higher employment probabilities for more recent birth cohorts of high school men in Canada at the same age compared with similarly educated men from earlier birth cohorts. This sensitivity indicates that it is the employment changes in the post 1989 period that are driving the employment cohort effects found in Figure 7. A likely reason is the very severe recession of the early 1990s. It may be that this recession had a differential impact on more recent birth cohorts than it did on earlier birth cohorts.

[21] We do not analyze the Australian data over this period because of the small number of surveys available.

This kind of excess sensitivity to business cycles for younger cohorts is often difficult to distinguish from permanent cohort differences when using a relatively short series of cross-sectional surveys. Rather than trying to fit a model that allows for this excess sensitivity to period differences by cohort, we note the sensitivity and allow for different interpretations of the cohort effects. A detailed investigation is left for future work.

The analysis is repeated for post-secondary men and university men in Canada. A similar pattern of earnings cohort effects to those of the second panel of Figure 4 is found. However, as was the case for the high school men, there is a reversal in the pattern of the employment cohort effects, implying higher employment probabilities for more recent birth cohorts than for earlier birth cohorts at the same age. For university men, the patterns are very similar to those found in Figure 4 and Figure 7, so that it appear that the sensitivity to the survey years is important for workers with education below the university degree level and that it is manifested primarily along the employment dimension.[22] We also carry out this sensitivity analysis for the U.S. men restricting the sample to the CPS surveys for 1982-1989. The results are very similar to what is presented in Figure 3 and Figure 6 and so are not reported.

This sample sensitivity provides one possible explanation for the difference in our results from those found by Card, Kramarz and Lemieux (1999). These data relate roughly

[22] As discussed by Beaudry and Green (2000), there was a change in the detail of education information in the SCF data from 1990 onward. Individuals who do earn a certificate from a program not requiring a high school diploma are categorized as having a high school education in the surveys prior to 1990 and a post-secondary education in the surveys starting in 1990. This does not affect the university degree composition but does imply a movement of relatively high skilled workers out of the high school group and into the post-secondary group after 1989. We investigate the importance of this by pooling the high school and post-secondary sample creating a non-degree sample. We then re-estimated all of the models for this non-degree case and found qualitatively similar results. In particular, when estimation is carried out over all years, non-degree men have strong earnings cohort effects and strong employment cohort effects. Also, when the sample is restricted to the 1982-89 period, the wage cohort effects remain but the employment cohort effects reverse

to the sample period covered by the two cross-sections of data used by Card, Kramarz and Lemieux (1999) for the United States and Canada. It may be that, for the sample period that they investigate, the deterioration in employment outcomes by birth cohort of less skilled men in Canada overall in the 1982-1996 period is not evident. That said, there are a number of important distinctions between our analysis and their analysis that could also explain the difference. The most important is the approach, since they employ instruments for changes in the level of demand and carry out a structural analysis to test for their importance. Our reduced form approach means that we are not identifying demand shifts from other possible causes of the change in the labour market setting that might impact on employment probabilities. Given this qualification, the sensitivity pattern that we find indicates that future research should investigate whether labour market inflexibilities in the post 1989 period led to the large employment birth cohort effects in Canada, rather than earnings adjustment which was the case for less skilled men in the United States.

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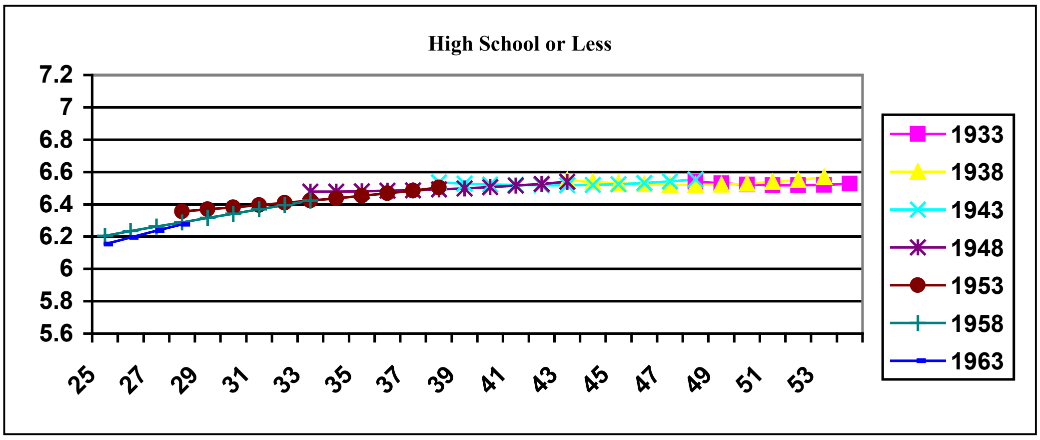
in a manner analogous to what is presented in the second panel of Figure 9. We conclude from this that this change in educational composition is not causing the sensitivity of the results to the choice of survey years.

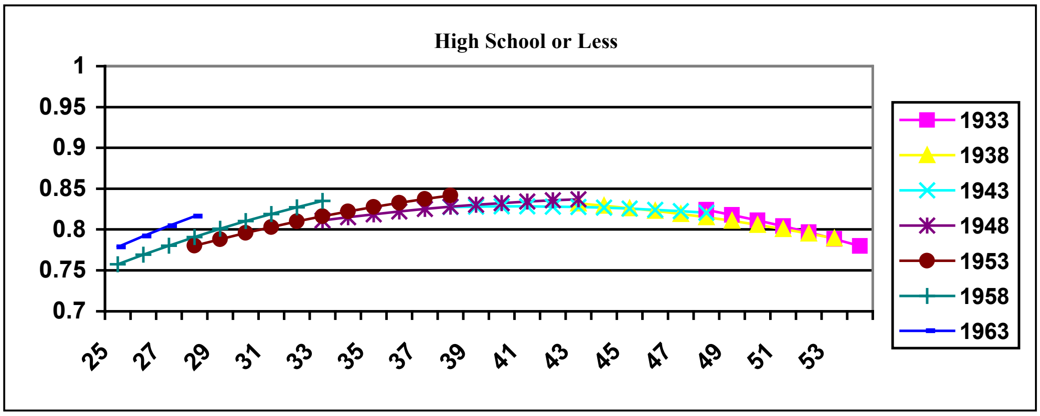
Figure 9

Earnings-Age profiles and Employment Probability-Age Profiles Estimated from 1982-89 Sub-Sample of High School Educated men in Canada

Earnings



Employment Probabilities



Note:

1. Profiles generated from earnings regression estimates and probit estimates over probability of employment versus non-employment.
2. The birth cohort years chosen are used as values for the continuous birth cohort variable in generating the predicted profiles.

8. Conclusions

In this paper, we analyse the earnings and employment probabilities of men by education level, birth cohort and age in the United States, Canada and Australia using a series of cross-sectional surveys for each country spanning the years 1982 through 1996. More recent birth cohorts of less-skilled men have experienced worse labour market outcomes than men from the same skill group but of earlier birth cohorts, ceteris paribus. In the United States, the deteriorating labour market outcomes appear as lower earnings but not lower employment probabilities. In Canada and Australia, the less skilled men from more recent birth cohorts experience lower employment probabilities as well as lower earnings and the magnitudes of the earnings declines across birth cohorts are smaller in magnitude than those found for the United States. This is consistent with the hypothesis that labour market institutions in Australia and Canada have prevented wage levels from declining in the face of negative demand shocks to the extent that they did in the United States, forcing employment to adjust as well.

We investigate the sensitivity of the results to the restriction of the sample years to 1982-1989 for Canada and the United States. The general pattern of the U.S. results is not sensitive to this restriction. However, for Canada, the employment cohort effects for the men without a university degree change direction implying higher employment probabilities for more recent birth cohorts of men, ceteris paribus. This sensitivity may be one explanation for the differences in our results compared with Card, Kramarz and Lemieux (1999) who do not find evidence that employment growth declined in response to adverse demand shocks in Canada. Determining the reason for the sensitivity of the results to the choice of survey year is an area of future research.

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Table 2
Coefficient Estimates:

OLS Earnings Models

Estimate	United States			Canada			Australia		
	H.S.	P.S.	Univ.	H.S.	P.S.	Univ.	H.S.	P.S.	Univ.
cohort	.0089 (0.50)	.0396 (1.55)	.0934 (3.79)	-.0399 (-2.24)	-.0042 (-0.16)	-.0594 (-1.76)	-.0201 (-0.38)	-.0508 (-1.04)	.1526 (1.70)
$(\text{cohort})^2$ /100	-.0204 (-2.02)	-.0310 (-2.14)	-.0507 (-3.63)	.0114 (1.12)	-.0088 (-0.58)	.0229 (1.17)	-.0044 (-0.14)	.0229 (0.82)	-.0845 (-1.63)
cohort× age/100	-.0635 (-0.32)	-.0523 (-1.83)	-.1125 (-4.09)	.0620 (3.03)	.0106 (0.36)	.0766 (2.01)	.0415 (0.70)	.0588 (1.06)	-.1894 (-1.84)
age	.1471 (6.31)	.3678 (10.7)	.4876 (14.9)	.1150 (4.11)	.2710 (6.24)	.3111 (5.49)	.0303 (0.38)	.0203 (0.27)	.3276 (2.41)
$(\text{age})^2/100$	-.3018 (-7.72)	-.0072 (-12.0)	-.8514 (-15.0)	-.3301 (-5.85)	-.6176 (-6.72)	-.7630 (-6.26)	-.1279 (-0.79)	-.0012 (-0.77)	-.0038 (-1.38)
$(\text{age})^3$ /1000	.0198 (6.17)	.0480 (9.74)	.0545 (11.7)	.0253 (5.36)	.0453 (5.79)	.0564 (5.50)	.0052 (0.71)	.0088 (0.70)	.0159 (0.68)
unemp. rate	-.0105 (-5.61)	-.0143 (-5.19)	-.0105 (-4.24)	-.7079 (-4.47)	-.6318 (-2.43)	-.5829 (-1.77)	.0147 (2.82)	.0079 (1.62)	.0042 (0.49)
intercept	4.152 (5.06)	.0841 (0.07)	-3.656 (-3.20)	6.091 (7.32)	3.066 (2.48)	3.990 (2.54)	6.801 (2.83)	7.759 (3.46)	-2.855 (-0.70)
p-value for overall F-test	.0001	.0001	.0001	.0002	.0001	.0001	.0001	.0001	.0001
Sample Size	192705	83035	104004	113326	48482	30132	7148	7172	2307

Note:
The value .0001 is used for all prob-values below .00015.

Employment Models

Estimate	United States			Canada			Australia		
	H.S.	P.S.	Univ.	H.S.	P.S.	Univ.	H.S.	P.S.	Univ.
cohort	.0874 (2.57)	-.0753 (-1.28)	-.0871 (-1.40)	-.0025 (-.062)	-.1151 (-1.66)	-.1832 (-1.50)	.1258 (0.61)	.2137 (0.80)	.0920 (0.15)
$(\text{cohort})^2$ /100	-.0295 (-1.52)	.0597 (1.78)	.0664 (1.87)	-.0052 (-0.23)	.0724 (1.82)	.1082 (1.54)	-.0762 (-0.63)	-.1307 (-0.83)	-.0259 (-0.07)
cohort× age/100	-.1657 (-4.37)	.0380 (0.58)	.0536 (0.78)	-.0233 (-0.51)	.0504 (0.66)	.1320 (0.97)	-.1762 (-0.75)	-.2661 (-0.87)	-.1827 (-0.26)
age	.1475 (3.34)	.1981 (2.59)	.2809 (3.53)	.0181 (0.31)	.1436 (1.35)	.2100 (1.18)	-.1160 (-0.47)	.0663 (0.21)	.4071 (0.56)
$(\text{age})^2/100$	.0366 (0.48)	-.4097 (-3.09)	-.5695 (-4.16)	.0745 (0.66)	-.3106 (-1.45)	-.5411 (-1.54)	.6203 (1.58)	.3698 (0.73)	-.4019 (-0.35)
$(\text{age})^3$ /1000	-.0213 (-3.43)	.0229 (2.08)	.0311 (2.74)	-.0167 (-1.78)	.0126 (0.69)	.0300 (1.01)	-.0648 (-2.05)	-.0529 (-1.30)	.0002 (.002)
unemploy. rate	-.0351 (-9.78)	-.0281 (-4.41)	-.0171 (-2.63)	-4.439 (-11.5)	-4.526 (-5.84)	-1.402 (-1.13)	.0060 (0.07)	.0727 (0.68)	-.2604 (-1.07)
intercept	-4.270 (-2.75)	.1050 (0.04)	-.7673 (-0.27)	1.162 (0.63)	3.035 (0.93)	3.956 (0.70)	-.9856 (-0.10)	-6.309 (-0.51)	-6.079 (-0.22)
p-value for overall F-test	.0001	.0001	.0001	.0001	.0001	.0001	.0001	.0001	.0023
Sample Size	243025	101724	125885	173079	62634	35927	10766	9997	2844

Note:
The value .0001 is used for all prob-values below .00015.

Table 3
P-values for Tests of Birth Cohort Equality from Regression and Probit Estimation

The Base specification is defined in equation (1) and includes a cubic in the age variable, a quadratic in the birth cohort variable and an interaction of the age and birth cohort variables as well as the aggregate unemployment rate. The Covariates specification includes the variable of the Base specification but also includes controls for region of residence, marital status, part time hours (only in the earnings regressions), main language spoken (in the Canadian data), and race (in the American data).

Two different birth cohort hypotheses are tested. The first is the `No Differences' Hypothesis and it restricts the coefficients on the linear cohort variable, quadratic cohort variable and the cohort/age interaction variable to all equal zero. The second is the `Fixed Differences' and it restricts the coefficient on the birth cohort/age interaction variable to equal zero. The test results are presented as p-values from the Likelihood Ratio tests.

High School or Less

Model	Specification	Cohort Hypothesis	United States	Canada	Australia
Earnings	Base	No Differences	.0001	.0001	.0001
		Fixed Differences	.6998	.0001	.4869
	Covariates	No Differences	.0001	.0001	.0001
		Fixed Differences	.0001	.0419	.0633
Employment Probability	Base	No Differences	.0001	.0001	.0021
		Fixed Differences	.0001	.4896	.4526
	Covariates	No Differences	.0001	.0001	.0259
		Fixed Differences	.0001	.1136	.9907

Post-secondary

Model	Specification	Hypothesis	United States	Canada	Australia
Earnings	Base	No Differences	.0001	.0001	.0015
		Fixed Differences	.0265	.6407	.2904
	Covariates	No Differences	.0001	.0001	.0065
		Fixed Differences	.5388	.0848	.0455
Employment Probability	Base	No Differences	.0001	.0001	.0349
		Fixed Differences	.4433	.3953	.3851
	Covariates	No Differences	.0001	.0001	.1125
		Fixed Differences	.0556	.2280	.3201

University Degree

Model	Specification	Hypothesis	United States	Canada	Australia
Earnings	Base	No Differences	.0001	.0001	.2005
		Fixed Differences	.0001	.0117	.0661
	Covariates	No Differences	.0001	.0001	.7628
		Fixed Differences	.4118	.0742	.4402
Employment Probability	Base	No Differences	.0001	.0001	.6377
		Fixed Differences	.0956	.1840	.7944
	Covariates	No Differences	.0001	.0001	.5052
		Fixed Differences	.0004	.5027	.8938

Note:
The value .0001 is used for all prob-values below .00015.

Appendix 1:

We construct the Australian minimum wage series as follows. According to the explanatory notes in Australian Bureau of Statistics 6312.0, there was a $10 safety net wage increase by the Australian Industrial Relations Commission in 1997 that raised minimum weekly wages – the lowest award wages - to $359.40. Prior to this, there were three $8 safety net wage adjustments between 1992 and 1997. These give us a minimum wage series for 1992 to 1997. We then compiled a minimum award wage index from Australian Bureau of Statistics data for female workers in the textile, clothing and footwear industry (comparable information for males was not available prior to 1985). We used this series as an approximation for wage changes that occurred to the lowest of the award wages. Based on wage levels reported in 1985, this industry has the lowest award wage level of any industry in Australia. We then started with the min wage in 1992 and used the percent movements in the award wage index taken from ABS 6312.0 to create a minimum wage series back to 1982.