Understanding the Wage Patterns of Canadian Less Skilled Workers: The Role of Implicit Contracts

David A. Green*and James Townsend†

Abstract

We examine the wage patterns of Canadian less skilled male workers over the last quarter century by organizing workers into job entry cohorts. We find entry wages for successive cohorts declined until 1997, and then began to recover. Wage profiles steepened for cohorts entering after 1997, but not for cohorts entering in the 1980s - a period when start wages were relatively high. We argue that these patterns are consistent with a model of implicit contracts with recontracting in which a worker's current wage is determined by the best labour market conditions experienced during the current job spell.

JEL codes: J31, O33

1 Introduction

It is, by now, well known that the real wages of workers with high school or lower education have declined substantially in the last thirty years. Beaudry and Green (2000) and Beach and Finnie (2004) show that, among males, birth cohorts of these relatively less skilled workers entering the Canadian labour force in the mid-1990s received weekly wages approximately 20% lower than those received by similarly educated cohorts entering the labour force in the late 1970s. Morissette and Johnson (2005) show a similar pattern in hourly wages for males with less than two years of job tenure without any substantial recovery up to 2004. Furthermore, these studies strongly suggest that recent cohorts will have lower real

*Department of Economics, University of British Columbia and Research Associate, Institute for Fiscal Studies, London

†Department of Economics, University of Winnipeg

lifetime earnings than previous cohorts, as their returns to experience and/or tenure are not higher than those of the preceding cohorts. These patterns are of concern because they affect approximately 35% of the Canadian workforce. Moreover, Friedman (2005) argues that in a period when the majority of workers are experiencing stagnation or decline in their wages, the openness and tolerance of a society tends to decline, suggesting that the difficulties of low skilled workers could be transmitted to the least fortunate in our society through political economy effects. For all of these reasons, the strong decline in real wages for the less educated is a matter of considerable concern.

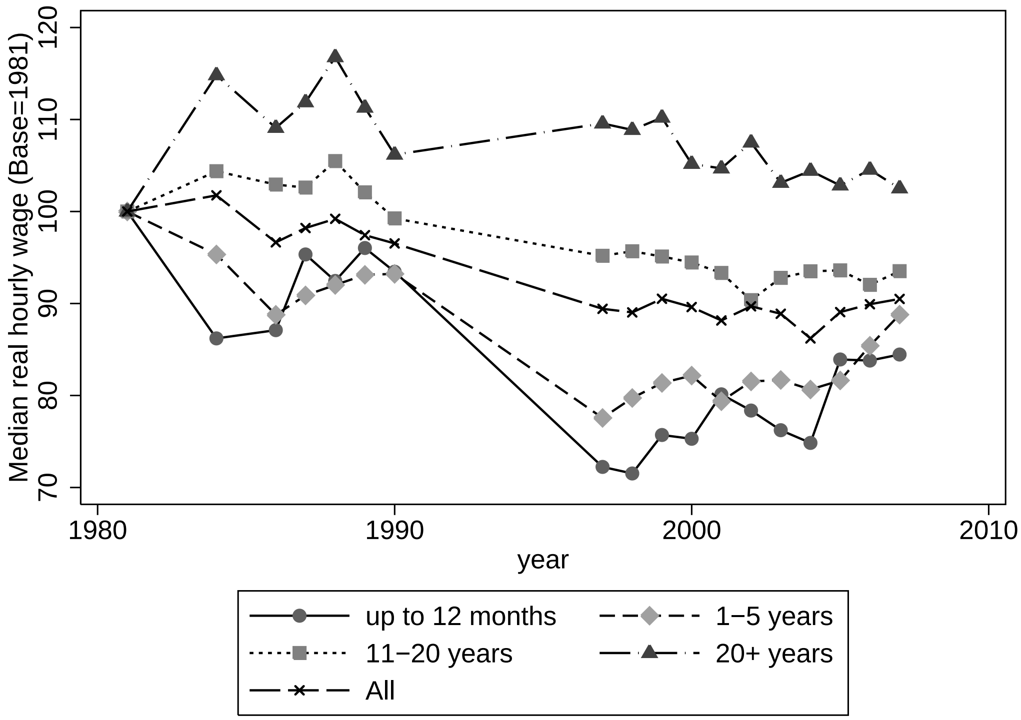
Our goal in this paper is to take a first step toward understanding the forces driving the wages for this large group. In particular, we are interested in whether there is a way to organize the divergent wage-tenure patterns of less skilled workers in order to put what needs to be explained in clearer focus. We argue that an implicit contract model with renegotiation provides an organizational framework that fits the data well. Because in that model, wages of higher tenure workers are related to those of entry workers through potential renegotiation, this implies that explanations for less-skilled wages should focus on patterns in entry wages. It also means that there is a certain amount of downward rigidity in low skilled wages that could become important as Canada enters tougher economic times.

To frame the specific questions we wish to pursue, it is useful to start with a simple description of real wage trends over the last quarter century.[1] Figure 1 shows median log real hourly earnings in paid jobs for men who have attended or completed high school by varying lengths of job tenure.[2] For all workers in the group, regardless of job tenure (indicated by the line labeled "All"), median real hourly earnings began to decline in the 1980s and continued to do so until the late 1990s. Over the entire period, the overall decline in the real median wage was 12%. For new job entrants (defined as those with up to 12 months of job tenure), the median real hourly wage plummeted during the recession in

[1] Throughout this paper we focus mainly on male wages because they are not affected by the potential selection effects accompanying the substantial rise in female labour force participation over our period of interest and, thus, we view them as providing a cleaner measure of the effects of underlying forces in the economy. The underlying data is described in detail in Section 3.

[2] The educational data is limited to two categories: 0-8 years and 9-12 years of schooling. We focus on the latter group, as this is the largest by far, making 75% of our sample in 1981 and increasing to 92% by the final year of our sample. Where applicable, corresponding figures and tables for men with 0-8 years of schooling are available in an online appendix.

Figure 1: Change in median log real earnings by length of job tenure, men, some or completed high school, Ages 20-64.



the early 1980s, partially recovered with the economic boom of the late 1980s, and then plummeted again with the recession in the early 1990s. In 1997, median hourly earnings for entrants were 26% less than they were in 1981. After 1997, entry wages increased, but by 2007 were still 15 % below their 1981 value. Wages for workers with 1-5 years of job tenure followed a similar pattern to those of new entrants, albeit with declines that were smaller in magnitude. In contrast, the median log real wage of workers with 11-20 years of job tenure was flat until 1997, at which point it began to decline slightly, while the wages for workers with longer job spells (20+ years) increased until 1997 and then began to decline gradually.[3]

Figure 1 points to a compression of the wage-tenure structure in the late 1980s and again during the strong labour markets after 1997. These were offset by expansions during poor labour market periods. Such a pattern does not fit well with the simplest spot market model

[3] In the online appendix, we present the corresponding figure for women with the same level educational attainment. We find a similar pattern, though both the decline and subsequent recovery of entry wages are smaller in magnitude.

in which all wages reflect a common productivity shock. However, it may fit with an implicit contract model in which firms offer workers insurance against adverse productivity shocks but are forced to match wage offers from other firms to retain workers when conditions improve (Malcomson, 1999). These models generate cohort effects based on time of job entry. When the economy is deteriorating, the wages of workers in ongoing job spells are insured and are based on conditions and information available at the time the job started, while those of new entrants reflect current conditions and only insure against further negative shocks. The result is an expansion in the cross-sectional wage-tenure profile. When conditions improve, wages in jobs that pay more than the current entry wage are unaffected, while wages in jobs that paid less than the current entry wage rise to match it. Thus, recontracting as conditions improve in the later part of the period under study would show up as increased returns to tenure for cohorts entering after 1995 and a compressed cross-sectional wage-tenure structure. Our primary goal in this paper is to investigate whether this type of model does indeed provide a good characterization of the patterns in Figure 1.

In addition to exploring the role of implicit contracts for explaining wage patterns in Canada, we make a methodological contribution to the literature on implicit contracts. Beaudry and Dinardo (1991) develop a model of such contracts over the business cycle, which has been tested using wage data from the U.S. (Beaudry and Dinardo, 1991; Grant, 2003), Canada (McDonald and Worswick, 1999) and the U.K. (Devereux and Hart, 2007) with varying degrees of success. As our wage patterns do not appear to be stationary in nature, we develop a methodology which allows us to test the recontracting model in the presence of a more general stochastic structure than the AR(1) process assumed by Beaudry and Dinardo.

In examining these patterns, we need to take account of strong compositional shifts occurring during this period. As Morissette et al. (2005) point out, unionization drops substantially over this period, particularly for new entrants, and the industrial affiliation of less skilled workers shifts as well.

The results of our investigation indicate that the recontracting model provides a reasonably good fit for the wages patterns of less skilled Canadian males over the last 25 years. These results are obtained after controlling for shifts in unionization and industrial employment, indicating that that the decline in entry wages depicted in Figure 1 occurred across industry-unionization groups. This suggests that demand for low-skilled workers decreased for much of the last quarter century, but that the wages of longer tenured workers were at least partly insured against this development by their employers.

In the very last section of the paper, we provide a brief examination of potential driving forces behind the movements in entry wages. We find that about one quarter of the decline in average entry wages for less skilled workers can be accounted for through the substantial declines in unionization and shifts in industrial structure in this period. Controlling for compositional shifts by focussing on nonunion wages, we find that about 60% of the decline in nonunion entry wages can be accounted for by movements in relative factor employment in a manner that fits with the type of induced technological change model described in Beaudry and Green (2003, 2005). But whatever the forces driving the entry wage, the results in this paper indicate that they are passed through the rest of the wage structure in such a way that they are felt for decades after.

The rest of this paper is organized as follows. In section 2, we outline the theory and empirical implications of implicit contract theories, and briefly explain how our empirical methodology follows from the theory. In section 3 we describe our data. In section 4, we present job cohort wage patterns and investigate whether they fit the implicit contract model from section 2. In section 5 we estimate our model separately for union and non-union workers to ensure that the patterns we discuss in section 4 are not an outcome of deunionization. In section 6, we briefly examine forces that could be driving the entry wage pattern. Finally, in section 7, we summarize our findings and make concluding remarks.

2 The Empirical Implications of Implicit Contract Theory

The properties of risk-sharing labour contracts are summarized in detail in surveys by Malcomson (1999) and Thomas and Worrall (2007). These models assume that productivity follows a stochastic process and that risk-adverse workers have limited access to capital markets. Workers are thus only able to insure against productivity shocks if firms offer insurance through a wage contract. The precise nature of the insurance depends on the extent to which the parties are able to commit to the risk-sharing contract. We outline two variants: 1) contracts that are binding on both parties (full commitment), and 2) contracts that are binding only on the firm (limited commitment with recontracting). Both these implicit contract models and a simple spot market in which all wages just equal a worker's current marginal revenue product have implications for how wages are related across job entry cohorts. In this section, we outline these models and their wage implications. Our presentation follows Malcomson (1999) and Beaudry and Dinardo (1991).

In the canonical implicit contract model, workers are homogenous and each firm hires a single worker. The output of a firm at time t is contingent on the state of the world, s_t, and is given by $\Phi(s_t)$. Let s^{t+j} denote the realization of a sequence of states $(s_t, s_{t+1}, .., s_{t+j})$, where the conditional probability of such a sequence occurring, contingent on beginning in state s_t, is given by $P(s^{t+j}|s_t)$ and s_t is observed by all agents. Firms are perfectly competitive and are assumed to be risk neutral. Workers are risk-adverse and care only about the wage in each period. Both the firm and the worker are infinitely lived and share a common discount factor β. Let $x(t)$ indicate the wage for new job entrants in period t and let $w(t', t)$ denote the wage in period t of an individual that started his current job in period t'. If wages are determined in the spot market, then the wage equals current output,

$$x(t) = w(t-j, t) = \Phi(s_t), \text{ for } j = 1, 2, ... \tag{1}$$

More specifically, workers from all cohorts defined by time of job entry receive a common wage equal to productivity in each period.

In the case where the contract is binding for both parties, the contract provides full insurance and competition drives expected profits of firms to zero. Since expected profits are zero, the wage is equal to the expected present value of productivity, *i.e.*, for a job starting at time t, the wage in period $t + j$ is given by

$$w(t, t+j) = x(t) = (1-\beta) \sum_{i=0}^{\infty} \beta^i \sum_{s^{t+i}} P(s^{t+i}|s_t)\Phi(s^{t+i}) \text{ for all } j. \tag{2}$$

In this case, the wage for each cohort is determined entirely by conditions at the time the job spell begins.

Finally, in the case where workers are fully mobile, firms will have an incentive to "bid away" workers working for other firms in some states. To eliminate such "bidding-away" opportunities, a contract formed at time t must satisfy the following inequalities:

$$\sum_{i=j}^{\infty} \beta^{i-j} \sum_{s^{t+i}} P(s^{t+i}|s_{t+j})[\Phi(s^{t+i}) - w_{t+i}(s^{t+i})] \leq 0 \tag{3}$$

for all $j = 1, ..., \infty$ and all realizations of s_{t+j},

where $w_{t+j}(s^{t+j})$ is the wage at time $t+j$ of a job beginning at time t, contingent on the realization of a sequence of states, $s^{t+j} = (s_t, s_{t+1}, ..., s_{t+j})$. The bidding-away constraint implies that j periods into the contract, the wage must be such that the expected profits of the firm are less than or equal to zero. The resulting wage contract is downward rigid: the wage is constant until a state s_{t+j} is reached where the wage for new contracts starting in this period, $x(s_{t+j})$, exceeds the current wage in the ongoing contract. To retain the current worker, the firm increases the wage in the ongoing job contract to match the wage being offered in new contracts. This implies that the wage ratchets up over time:

$$w(t, t+j) = w_{t+j}(s^{t+j}) = max\{x(s_t), ..., x(s_{t+j})\}. \tag{4}$$

Let Ω^{t+i} be the set of all sequences of states, s^{t+i} up to time $t+i$ for which it has not become profitable to bid away a worker who accepted a starting wage $x(s_t)$ at time t. Equation (3) implies that the expected profits at the time the wage was offered are zero. It is straightforward to show that the offered wage is,

$$x(s_t) = \frac{\Phi(s_t) + \sum_{i=1}^{\infty} \beta^i \sum_{s^{t+i} \in \Omega^{t+1}} P(s^{t+i}|s_t)\Phi(s^{t+i})}{1 + \sum_{i=1}^{\infty} \beta^i \sum_{s^{t+i} \in \Omega^{t+1}} P(s^{t+i}|s_t)} \tag{5}$$

That is, the wage equals the expected present value of productivity, conditional on productivity not rising enough at some future date such that it is profitable for another firm to bid the worker away at a higher wage rate. In other words, the implicit contract insures the worker against adverse productivity shocks but does not involve income smoothing when productivity rises, as it is not possible to prevent the worker from leaving for a higher wage elsewhere. Note that the offered starting wage will fall from one period to the next if productivity falls across periods. The starting wage will also fall if the magnitude and/or probability of adverse shocks increases, as workers pay the actuarially fair price for insurance against these shocks in the first period.

These three models can be nested in a single equation. For a job beginning in period t, the wage at period $t+j$ is given by

$$w(t,t+j) = \begin{cases} x(s_{t+j}) & \text{spot market model} \\ x(s_t) & \text{contracts with perfect commitment} \\ max\{x(s_t), x(s_{t+1}), .., x(s_{t+j})\} & \text{contracts with worker mobility.} \end{cases} \tag{6}$$

As part of our our empirical assessment of these models, it is convenient to define several variables. Let $Change(t,t+j)$ denote the change in the entry wage between the time a job starts, t, and a later period, $t+j$:

$$Change(t,t+j) = x(s_{t+j}) - x(s_t). \tag{7}$$

Let $Ratchet(t,t+j)$ be a variable that measures the largest increase in the start wage between period t and period $t+j$. If no positive increase occurs in this interval, $Ratchet(t,t+j)$ takes a value of zero:

$$Ratchet(t,t+j) = max\{0, Change(t,t+1), .., Change(t,t+j)\}. \tag{8}$$

Thus, $Ratchet$ takes a value of zero in the first year for a job cohort and remains at zero until reaching a year $t+k$ where $Change(t,t+k) > 0$. It then takes on and remains at this value until reaching a later year, k' where $Change(t,t+k') > Change(t,t+k), k' > k$, at which point it ratchets up to the new higher value of $Change(t,t+k')$. If subsequent entry wages between periods $t+k'$ and $t+j$ are lower than $x(t,t+k')$ then the value of $Ratchet$ remains at $Change(t,t+k')$.

Let $Below(t,t+j)$ be the difference between $Change(t,t+j)$ and $Ratchet(t,t+j)$:

$$Below(t, t+j) = Change(t, t+j) - Ratchet(t, t+j). \tag{9}$$

This variable measures how far the current start wage is below the highest start wage experienced so far in the job spell. Using these variable definitions, the following equation nests all three models:

$$w(t, t+j) = x(t) + \beta_R Ratchet(t, t+j) + \gamma Below(t, t+j), \tag{10}$$

$$\begin{cases} \beta_R = 1, \gamma = 1 & \text{spot market model.} \\ \beta_R = 0, \gamma = 0 & \text{contracts with perfect commitment.} \\ \beta_R = 1, \gamma = 0 & \text{contracts with worker mobility.} \end{cases}$$

where β_R measures how responsive the wages in an ongoing spell are to the highest entry wage experienced during the job spell, and γ measures the responsiveness of the wage to movements in the current period entry wages that are below the highest entry wage experienced to date during the job spell.

We assume that the wage determination in the models set out earlier occur within labour markets defined by skill and province. To incorporate this element, we write individual i's wage as a function of observed and unobserved characteristics as well as of the factors set out in equation (10):

$$ln\, w_{t,t+j,i} = \alpha(t) + X_{i,t+j}\beta + \beta_R Ratchet(t, t+j) + \gamma Below(t, t+j) + e_{i,t,t+j} \tag{11}$$

where $X_{i,t+j}$ is the vector of observable individual and job characteristics of individual i observed at time $t+j$, $e_{i,t,t+j}$ corresponds to unobserved characteristics, $\alpha(t)$ is a fixed effect for the entry cohort beginning jobs in period t, and $Below(t, t+j)$ and $Ratchet(t, t+j)$ are measured as differences in the logged entry wage. We use the log wage as the dependent variable in order to provide comparable estimates to standard human capital regression results.

The business cycle variants of the above model that Beaudry and Dinardo (1991) develop are special cases of the models nested in equation (11). By assuming that productivity follows an AR(1) process, and explicitly modeling the reservation wage of the marginal worker, Beaudry and Dinardo are able to replace the entry wage in a given period with the unemployment rate, which is inversely related to the entry wage in all three models. The spot market implies that wages are negatively correlated with the current unemployment rate; the perfect commitment model implies that wages are negatively correlated with the unemployment rate at the time the job starts; and the worker mobility model implies that wages are negatively correlated with the lowest unemployment rate experienced during a job spell. The advantage of making these assumptions is that start wages do not not need to be measured directly. For the US in the period that Beaudry and DiNardo study, their assumption of a stationary process for productivity appears to be appropriate, but it seems unlikely to hold for the less-skilled in Canada in our period, where there has been a long downward trend in wages. In that case, the unemployment rate does not provide a good measure of recontracting pressures. In particular, the entry wages of workers entering in the recession of the early 1980s could be above those of job starters in the strong labour market of the late 1990s. In that situation, the early 1980s job starters should not seek to renegotiate their wages in the late 1990s even though the unemployment rate was much lower than when they started.

In our empirical implementation, we also allow for upward sloping wage-tenure profiles. The implication of those profiles for how we should think about the cohort/contracting effects depends on the source of the positive tenure effect. One possibility is that the slope reflects worker and firm sharing of investments in and returns to specific human capital (Hashimoto, 1981). If we assume the impact of such investments on productivity are perfectly foreseeable then we can potentially treat them as separable from the uncertain (macro) productivity shocks set out in the model. Workers, of course, might prefer to have their portion of the specific human capital returns delivered to them as a constant wage but firms may prefer an upward sloping profile to address potential shirking. In either case, because a new firm could not capture the benefits of the specific human capital investment, the relevant comparison in any renegotiation would be between the going entry wage (without tenure) and the worker's current wage inclusive of tenure effects. This raises the possibility that a given cohort - if it is far enough along its profile - will not have a renegotiation of its wage even if the going entry wage is above the entry wage for the cohort. We discuss and check the specific empirical implications of this version of the model in section 4.3.2.

A second possibility is that a positive wage-tenure profile reflects returns to general human capital. In that case, outside firms interested in bidding a worker away will offer a worker a wage inclusive of the returns to job tenure. The current firm will not offer to smooth the wage over future gains coming from human capital accumulation, as this involves offering higher wages in the current period in exchange for lower wages in the future. Since the worker is more productive in the future, other firms would have an incentive to bid this worker away with a higher wage, even in the absence of productivity shocks. However, the current firm may still insure the wage profile against stochastic shocks. Interestingly, the implied empirical specification in this case is still based on the entry wage, as in (11). To see this, note that an outside firm will offer the worker the entry wage plus the value of the human capital reflected in the wage-tenure profile. But since the current and outside firms place equal value on that human capital, the comparison again comes down to a comparison of entry wages. While they do not explicitly discuss the source of any wage-tenure profile, this is the approach to tenure implicitly adopted in all previous papers on wage contracting (Beaudry and Dinardo, 1991; McDonald and Worswick, 1999; Grant, 2003; Devereux and Hart, 2007).

Finally, we have presented our discussion in terms of productivity and human capital investment but we could also present the discussion in terms of matching models. In particular, in matching models, if match quality is revealed gradually and poor matches are terminated, wages (which equal the average productivity of the workers remaining with the firm) will rise with tenure. An overall decline in productivity could then be modeled as a decline in the mean of the match quality distribution. If this affects ongoing and new matches equally and there is no change in the costs of forming or dissolving a match then there is no reason to believe that the selection associated with the match process will change over time, even as the trend changes. If, instead, the cost to a firm of dissolving a match rises in a tight labour market because the time their capital will be idle until a new match is found lengthens then selectivity could differ over our time period. In particular, in the tighter labour markets after 1997, matches would be less likely to be dissolved which would imply both lower average productivity for a given cohort and a lower slope to the wage-tenure profile. Thus, such selectivity, if it exists, will offset to some extent the tendency toward higher and steeper profiles that we have discussed.

3 Data

The data behind Figure 1, and the data we will use in our investigations, come from a series of datasets, all of which are based on the common sampling frame provided by the Labour Force Survey (LFS). The LFS is a large, national survey used to calculate labour force statistics but, periodically, survey respondents are asked also to respond to special surveys, and it is the latter we use here. More specifically, we use the 1981 Annual Work Patterns Survey (AWPS), the 1984 Survey of Union Membership (SUM), the 1986-87 and 1988-90 Labour Market Activity Surveys (LMAS), the 1995 Survey of Work Arrangements (SWA) and the LFS itself for the years 1997 through 2007.[4] These surveys were chosen because they allow for calculation of consistent measures of the hourly wage of jobs and because they include a consistent question on job tenure (asking respondents when they started their current job in the AWPS and LFS and when they started a given job in the LMAS). The LMAS are two panel datasets but we use them as cross-sections with the weights provided for that purpose. We limit our analysis to individuals between the ages of 20 and 64. For workers between the ages of 15-19, a large proportion of job entrants have wages at the provincial minimum, suggesting that minimum wage changes explain much of the wage variation for this group. The fraction of workers receiving wages at or near the minimum wage declines substantially for older workers, suggesting that the minimum wage is not the main driving force of the patterns we outline.

Data comparability is clearly an important issue in using a series of different datasets over an extended period of time. Morissette and Johnson (2005) use some of the same

[4]The LFS did not collect wage data prior to 1997.

data for an examination of job quality and wage trends and include a discussion of the comparability of the hourly wages constructed in each. They raise two issues of importance. First, the specific questions on earnings in all years include the request to report "usual wage or salary before taxes and other deductions," but the LMAS after 1986 and the LFS tell the respondents to include tips and commissions in their earnings while the AWPS and the first year of the LMAS make no explicit reference to tips and commissions. Further, the 1987-1990 versions of the LMAS ask respondents to include overtime while the other surveys make no mention of overtime earnings. Similarly, the surveys for our two earliest years do not ask respondents about overtime when asking about hours worked while all the remaining surveys ask respondents to omit overtime. Given that wages are constructed by dividing earnings by hours of work, these two biases work in the same direction. The differences in earnings questions will tend to lead to respondents reporting lower earnings in the earliest surveys than the later ones, and the hours questions will tend to lead to higher reported hours in the earlier surveys. Combined, these lead to lower calculated hourly earnings in the initial two datasets, implying that the negative trend portrayed in this data is, if anything, an understatement of the actual trend.

The second issue has to do with imputation when respondents do not report their earnings. Morissette and Johnson (2005) report that both the LMAS and the LFS use class of worker, province, gender, age group and education level as part of their imputation. The LFS, in addition, uses student status, a renter/owner indicator, and occupation. The LMAS uses union status while the LFS does not. As shown in Hirsch and Schumacher (2004), differences in imputation can have noticeable impacts on wage differentials. In our case, moving from imputations that include union status to ones that do not means that union workers are imputed to have lower wages than they should, with the opposite being true for non-union workers. Unfortunately, Statistics Canada does not include imputation indicators on any of its datasets so all we can do is keep this issue in mind.[5]

Shifts in industrial composition over time are of considerable interest in the analysis that follows and, as a result, issues relating to the comparability of industry categories in the

[5]Using the Survey of Consumer Finance (SCF), which uses a consistent measure of earnings across the years it covers, we find a very similar patterns for hourly earnings over the period 1981–1997 to those depicted in Figure 1. Details of this exercise are available in the online appendix.

data must also be addressed. In 1997, the Standard Industrial Classification (SIC) scheme was replaced with the North American Industrial Classification Scheme (NAICS). Although Statistics Canada released revised data for LFS incorporating the NAICS, the supplemental surveys were not revised. We use an earlier version of the 1997 LFS, which reported industry of affiliation using both the NAICS and the SIC, to produce a concordance between the two systems. Using this concordance, we construct 14 industry categories based on the NAICS.

4 Job Cohort Specifications

4.1 Basic Data Patterns

We turn, now, to investigating the implications from the discussion in section 2. A key feature of the datasets we use is that each reports job tenure as a continuous variable. This allows us to construct and follow cohorts defined by year of job entry, which fits with the implicit contract models discussed earlier. Thus, a cohort will consist of a set of people who all started their current job in the same year, have the same gender and fall in the same education category. Given the key role that entry wages play in the models that we are interested in, we only retain cohorts for which we are able to observe wages at the start of the job spell. Our main focus is on men with 9-12 years of schooling. We emphasize this group because: 1) job entrants in this group experienced large wage declines over our study period, as documented in Section 1, and 2) this group is heavily represented in our sample, allowing for reliable estimates of the entry wage in each year of our sample.[6]

We do not have true panel data to follow these cohorts but we can construct synthetic cohorts. In the case of the first (1981) cohort, for example, we can construct their average wage in 1986 (from the LMAS data) by calculating the average wage for high school educated men who have 5 years of job tenure. For synthetic cohorts to provide consistent estimates of the true cohort profile, we require that the population from which the synthetic cohort is drawn at each point in time does not change. This is less of a problem when constructing job start cohorts than is the case in the more common birth cohort approach. For example,

[6]Men with 0-8 years of schooling experience similar wage patterns. We document this in the web appendix. However, over our sample period, this group makes up a small and declining fraction of men with less than 12 years of schooling. In 1981, 25% of men fell into this group. By 2007 the corresponding figure is 8%.

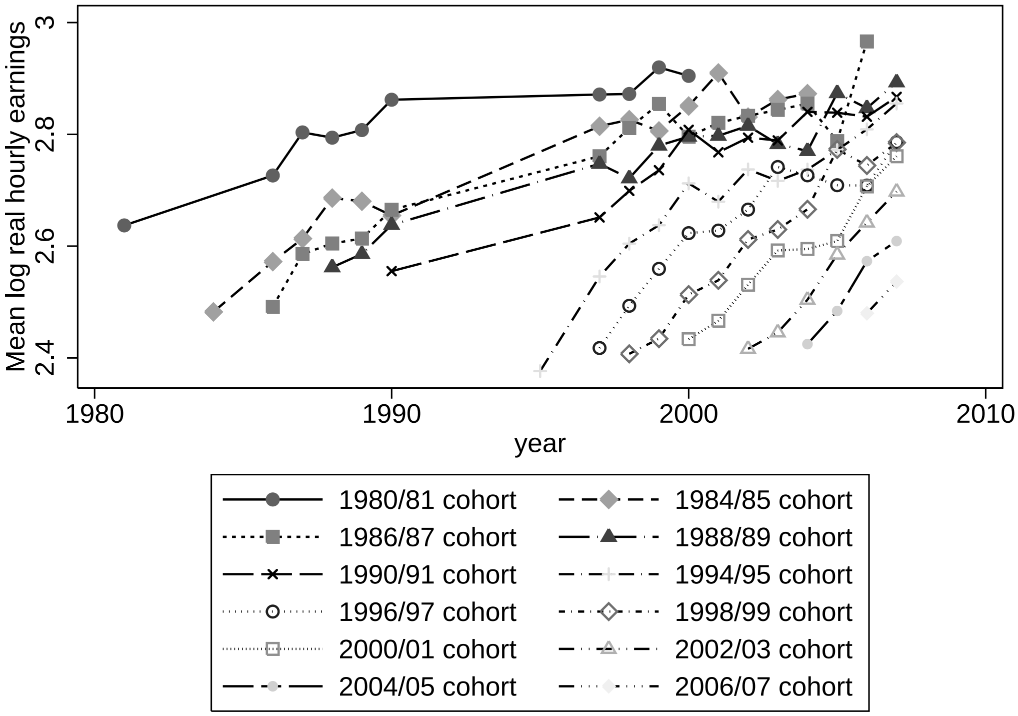
with birth cohorts, one might be concerned that immigrants enter the population, changing the composition of a given birth cohort between one cross-sectional dataset and the next. In our case, this is not possible since immigrants could not enter the economy with positive years of job tenure already assigned. The number of observations at varying lengths of job tenure is reported for each cohort in Appendix A.

In Figure 2, we present the raw data: average wages assembled by job entry cohort.[7] Several patterns stand out. First, the profiles show a pattern of rising substantially over the first decade or so of a job and then flattening out. Second, the earlier cohorts have higher starting wages, and higher profiles in general, than the later cohorts. Third, there appears to be a break in the data beginning with cohorts starting jobs approximately in the mid-1990s. Before that point, the cohort profiles tend to shift down in parallel, though with relatively clear effects of the late-1980's boom in the opposite direction. After about 1996, the profiles become steeper and entry wages rise. Notice that the mean wages at entry follow a pattern much like that of the median entry wages in Figure 1.

The lines are difficult to disentangle in parts of Figure 2. To provide a simpler picture of the cohort patterns, we regress wages on a complete set of cohort dummies and a complete set of interactions of those dummies with a spline in years of job tenure, thus creating a set of smoothed, cohort specific wage-tenure profiles. The spline allows for a different slope after 10 years of tenure and is tied at the 10 year point. We plot those profiles for a selection of cohorts in Figure 3. It is even easier to see both the break in the data with the mid-1990s cohorts and the pattern of entry wages described for Figure 2. It is again interesting to note the relatively parallel nature of the profiles for the pre-1997 cohorts. In particular, there does not appear to be strong evidence that these earlier cohorts shared in the steepening of the tenure profile witnessed for the post-1997 cohorts. The 1984/85 cohort, which entered the market during a recession, differs somewhat from the other 1980s cohorts. In the context of a model of implicit contracts with mobility, this would be the pre-1997 cohort most likely to experience wage gains as the economy improved in the late 1980s.

[7] In Figures 2 and 3, we define cohorts by two year groups in order to provide a smoother depiction of trends. However, cohorts are defined by single entry years in the econometric work that follows.

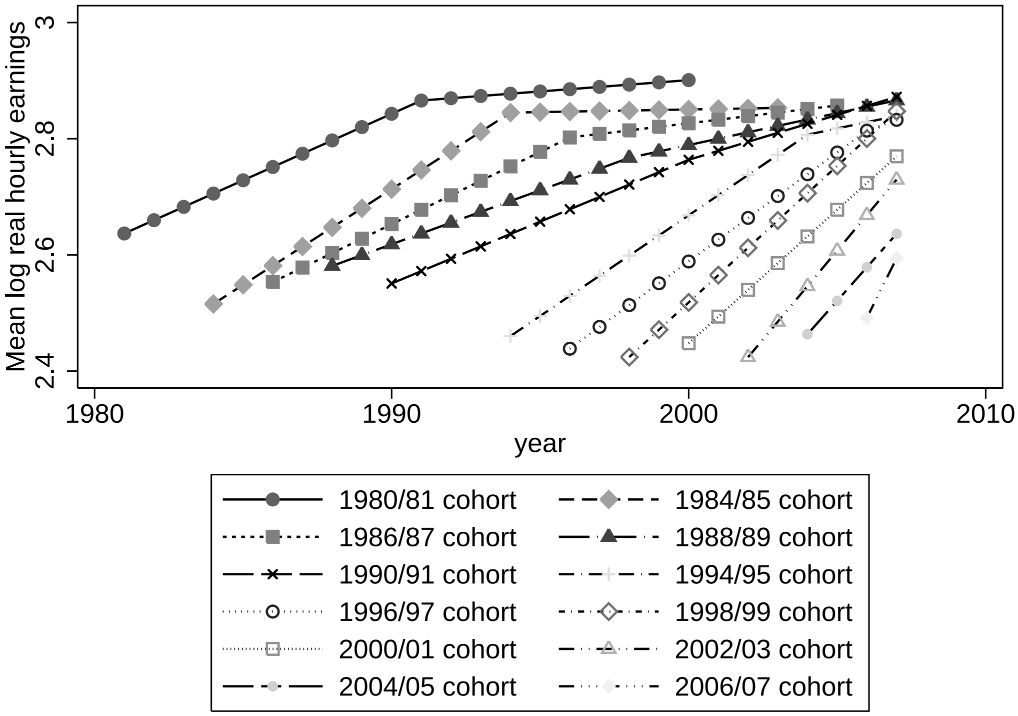
Figure 2: Mean Hourly Earnings, LFS, High School Males, By Job Entry Cohorts



It is possible, with this data, to track job start cohorts for different age groups. Thus, we can see whether the patterns in entry wages depicted in Figures 2 and 3 are present for all age groups or whether the declines across cohorts we have observed so far are just disguised birth cohort effects arising because new job starters are disproportionately young. If these patterns really reflect birth cohort effects (arising, perhaps, because of differences across cohorts in cohort size or the quality of schooling they faced) then we would expect not to see the long term declines in entry wages for older new job starters. Figure 4 contains a plot of entry wages (wages for workers with up to 1 year of job tenure) for various age groups for the years covered by our data.[8] These data show a clear impact of age on earnings (which can be interpreted as a standard experience effect given that we are controlling for schooling), with 45-54 year olds having wages that are approximately 40% higher than those for 20 to 24 year olds in 2007. The time paths of entry wages are also remarkably

[8]To create a clean presentation, we exclude 55 to 64 year old entrants from Figure 4. The wage patterns are considerably noisier for this group, which may reflect either the small number of entrants at these ages (less than 5% of job entrants in any year are over 55) or the different types of jobs taken after early retirement.

Figure 3: Fitted Mean Hourly Earnings, LFS, High School Males, By Job Entry Cohorts

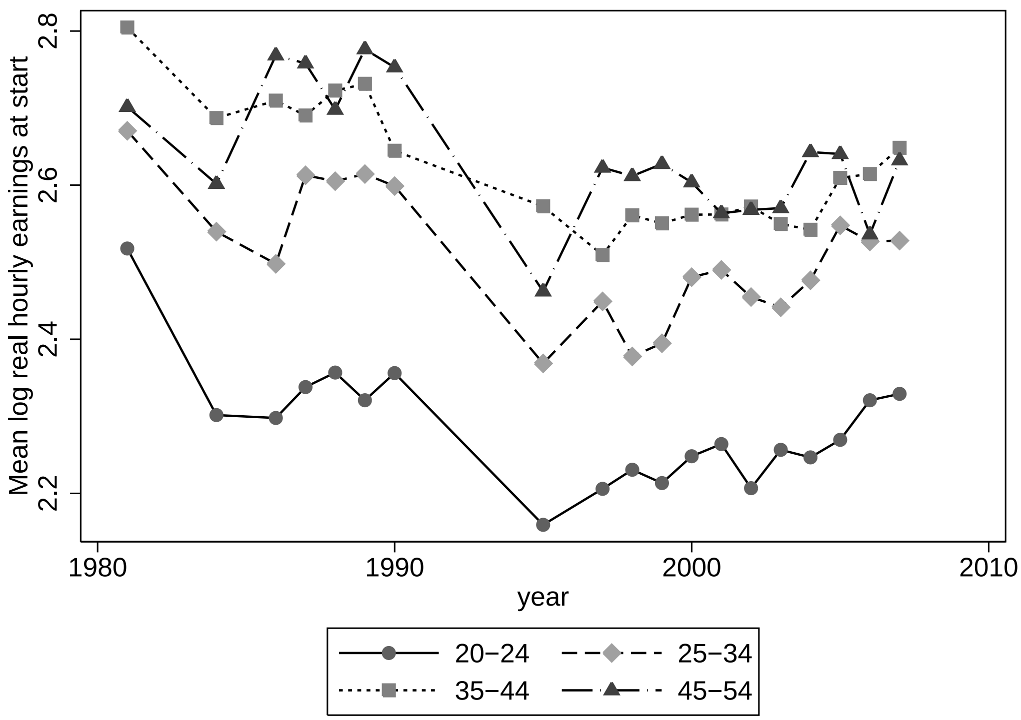


similar across the age groups, showing the same pattern of strong decline from 1981 to 1997 followed by either stagnation or slight recovery.[9] Given this, there is little evidence that the patterns we are picking up are masked age (or birth cohort) effects. Instead, it appears that a simple specification with constant age premia and a common trend in starting wages for all age groups characterizes the data quite well and this is the specification we use in estimating equation (11). This points to implicit contract models rather than models of forces affecting birth cohorts as the best way to approach the cohort patterns set out in Figures 2 and 3.

These simple data plots provide initial evidence that the type of implicit contracting model with re–negotiation we described in section 2 provides a useful framework for thinking about shifts in the Canadian wage structure for less educated workers, assuming a U-shaped trend for productivity. In particular, the 1981 cohort would have entered prior to the

[9]The largest group of job starters in any given year is the 20-24 grouping. Using the time series underlying Figure 4, the correlations between the start wages for this age group and other age groups are listed in parentheses: 25-34 (.89), 35-44 (.85), 45-54 (.59), and 55-64 (.40).

Figure 4: Mean Log Hourly Earnings, High School Men, Up To One Year of Job Tenure, By Age



recession of the 1980s when productivity was high and the subsequent secular decline in wages from 1981–97 may have been viewed as a low probability event, making the cost of insuring against it low. This would result in high entry wages. After the recession, wages did not fully recover, potentially reflecting a realization among economic agents that productivity for this group of workers had fallen and was likely to continue to fall. Thus, lower productivity and a higher cost of insurance against further negative shocks would have continued to erode entry wages. After 1997, entry wages leveled off and began to increase. In the context of the model, this would reflect a reversal in the productivity trend. The rising entry wages, in turn, would allow relatively recently hired workers to recontract. However, workers hired during the 1980s would still be on high wage paths that would imply no renegotiation. Thus, wage profiles for post-1995 entry cohorts would steepen while those for the 1980s cohorts would appear unaffected by the turnaround. This is what we see in the data plotted by cohorts and accounts for the pattern in Figure 1, where increases in cross-sectional wage-tenure differences in the 1980s were replaced by decreases in those differences after 1997. In contrast, there is no evidence of the type of pattern one would predict from a simple spot market model; a pattern in which real wages decline over time within specific job cohorts during the early 1980s and early 1990s.

4.2 Econometric Specifications

We now turn to a more careful, econometric examination of the wage patterns seen in the previous figures. We estimate equation (11) using a two-stage procedure. In the first stage, we regress the log of individual wages on a vector of individual and job characteristics, X_{it}, and a complete set of start year-job tenure interactions, $\pi(t, t+j)$. The estimated equation is thus of the form:

$$ln\, w_{ijt} = X_{ijt}\beta + \pi(t, t+j) + e(i, t, t+j), \tag{12}$$

where, as before, i denotes an individual, t denotes the year in which the current job started, j indicates years of job tenure in the current job, $t+j$ indicates the year in which the individual is observed, and X is a vector of individual characteristics which we describe in the next section. In the second stage, we estimate the following model:

$$\pi(t, t+j) = \alpha(t) + g(j, t) + \beta_R Ratchet(t, t+j) + \gamma \cdot Below(t, t+j) + \nu(t, t+j) \tag{13}$$

where $\alpha(t)$ is a cohort effect for the cohort entering jobs at time t, $g(j, t)$ is a cohort-specific wage-tenure profile for individuals that started jobs in year t with j years of job tenure, and the remaining variables are those described in section 2.

We estimate Equation (12) using the sample weights accompanying the various data sets. As the dependent variable in equation (13) is obtained by estimating equation (12), we estimate this equation using WLS with the inverse of the standard errors of the estimates used as weights. The two-stage procedure corrects the standard errors of β_R and γ for correlations across the error terms, $e(i, t, t+j)$, of individuals that started jobs at the same time and are observed in the same year.

Based on observations from Figure 2, we initially adopt a tied spline specification for the wage-tenure profiles, $g(j, t)$, in which the wage profile for any cohort is allowed to have different slopes before and after 10 years of tenure. As our *Ratchet* variable is based on changes in the start wage relative to the start wage at the time a job spell began, we only work with those cohorts for which we observe a wage at the time of job start. Using the data described so far, this includes the 1981, 1984, 1986–90, 1995 and 1997–2007 cohorts, giving us 19 cohorts. The 1984 SUM and the 1995 SWP are offshoots of the LFS which include only categorical variables for job tenure. As a result, we only use these data sets to provide observations for job starters in these years, since the categories allow us to identify workers with at most one year of job tenure.

4.3 Results

4.3.1 Compositional Effects

Table 1 presents information on the job characteristics of job starters in the years 1981, 1984, 1989, 1997 and 2007. The years 1981, 1989 and 1997 were chosen because they occur at similar points in the business cycle. There has been a marked decline in the percentage of new job starts involving union membership (a point also made in Morissette and Johnson (2005)); this development began in the 1980s and accelerated in the 1990s. In 1981, 35.5% of job starters were members of unions; this number declined to 26.7% by 1989 and plummeted to 14.0% by 1997. As the figures from 2007 indicate, unionization rates of starters remained low throughout the last decade of our sample. There have also been notable changes in the industrial composition of new jobs for high school men. In 1981, nearly a quarter of new jobs were in manufacturing. By 2007, that number had fallen to 14.7%. The proportion of jobs in the primary sector also declined. Declines in these industries were offset by increases in the various service sectors. The age distribution of starters changed over the study period: in later years, new starters were more likely to be over the age of 35. 1984 is one of the few years in our data set which corresponds to the trough of a business cycle. The most striking difference between this year and 1981 is the sharp decline in jobs in construction.

Given the substantial compositional shifts depicted in Table 1, we are interested in how

Table 1: Job characteristics of male job starters, high school educated, selected years

	Start Year				
	1981	1984	1989	1997	2007
Age					
20-24	36.2	36.1	30.9	33.1	33.2
25-34	36.7	35.0	40.8	32.2	27.7
35-44	14.0	17.2	16.2	21.5	20.1
45-54	8.6	8.5	8.6	9.4	13.5
55-64	4.6	3.2	3.6	3.8	5.6
Union	35.5	25.6	26.7	14.0	13.8
Agriculture	1.5	1.5	2.0	2.2	1.4
Primary	7.7	5.8	5.1	5.8	4.8
Construction	19.1	15.1	21.4	16.8	19.1
Manufacturing	23.7	23.2	22.5	19.3	14.7
Trade	17.1	22.6	19.1	18.1	18.2
Utilities	0.9	0.8	0.6	0.2	0.4
Transportation & Storage	9.4	8.3	9.0	9.3	9.7
FIRE	1.8	2.4	1.5	1.9	3.4
Education	1.4	1.2	1.7	1.0	1.6
Public admin.	3.5	2.7	3.1	1.7	1.6
Accommodations/Food	6.0	5.6	5.2	8.3	7.2
Health and Soc. Assist.	1.5	1.8	1.3	1.1	0.8
Professional Services	2.7	4.2	2.8	6.7	10.6
Other Services	4.0	4.8	4.6	7.6	6.6
Mean Log Wage	2.64	2.48	2.55	2.40	2.50

Note: Job starters are defined as workers ages 20-64 with up to one year of job tenure. All means are generating using the sample weights included with the respective surveys.

much of the overall wage patterns in Figures 1-3 can be explained by those shifts. We can provide an answer using the first stage wage regression described in the previous section. In particular, the regression includes indicators for the age and province of residence of the individual, controls for industry and union status of the individual's main job, and the complete set of tenure-year interactions. The results yield expected patterns, such as a roughly quadratic shaped wage-age profile and a positive union wage premium.[10]

Figure 5: Entry Wages for High School Men, Relative to 1981

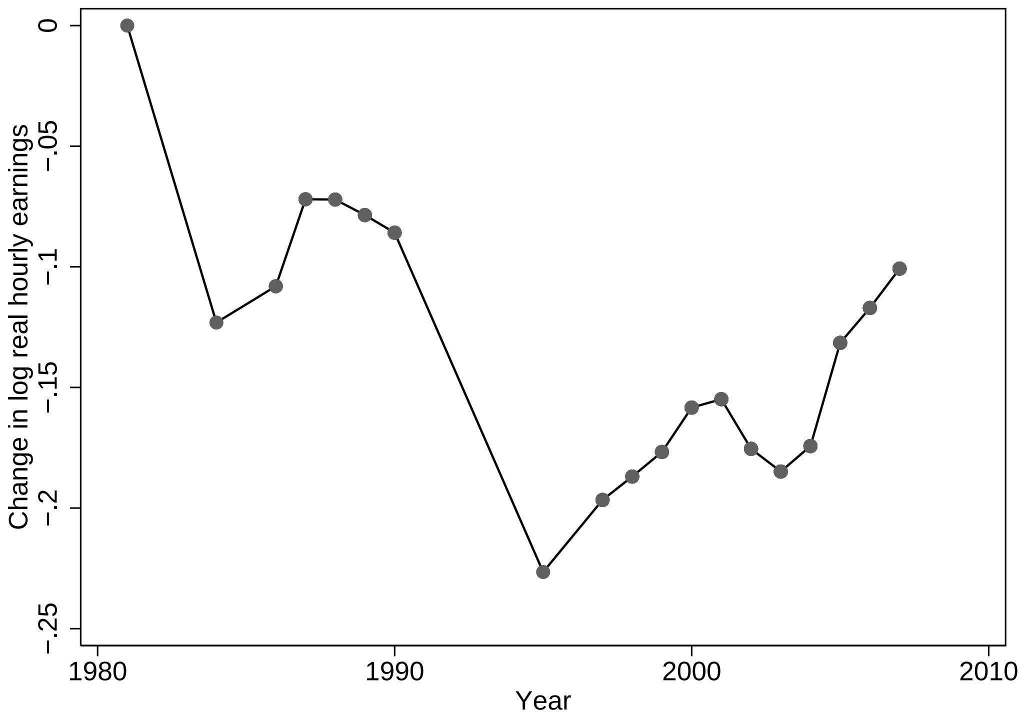


Figure 5 shows the estimated cohort effect at the time of job start, $\pi(t, t)$, relative to that of the 1981 cohort, using the first-stage specification reported above. These effects are measures of the entry wage in each year, after controlling for differences in the characteristics and jobs of entrants across years. The pattern is similar to what is shown in the previous figures. Entry wages declined substantially in the early 1980s, partly recovered in the late 1980s, fell again in the early 1990s, and then followed a general pattern of recovery for the remainder of the period. As age, industry and union controls are included in the first stage, these results indicate that the pattern for entry wages that we observe are being driven

[10]The estimates of the first stage coefficients are reported in the web appendix.

Table 2: The Roles of Unionization and Industrial Patterns in Accounting for Entry Wage Patterns

	1989-81	1997-1981	2007-1981	2007-1997
Unionization	-.02	-.04	-.04	.00
Industry	.00	-.02	-.02	.00
Explained	-.02	-.06	-.06	.00
Total	-.09	-.24	-.13	.10

Note: Decompositions are based on average union and industry effects over the entire study period and are expressed in differences in the mean log hourly wage, after controlling for age and province differences across entry cohorts.

largely by forces other than compositional shifts. Morissette and Johnson (2005) reach the same conclusion for the period 1981–1997 using a similar methodology.

To provide a more concrete measure of the role of compositional shifts, in Table 2, we present the results of a simple decomposition that focuses on the roles of deunionization and industrial shifts using the estimates from the first stage regression. As shown in Table 1, the fraction of job starters that were members of a union in their job fell from 35.5% in 1981 to 26.7% in 1989, and to 14.0% in 1997. After 1997, unionization rates remained flat. Based on a average union premium of .171 log points, this accounts for roughly .02 log points of the wage decline from 1981 to 1989 and .04 log points of the decline from 1981 to 1997. There was also a shift from jobs in manufacturing and the primary sector towards jobs in services, most of which took place after 1997. Based on estimates of the various industry premia over the period, shifts in industrial patterns account for none of the difference in entry wages between 1981 and 1989, and .02 log points of the decline between 1989 and 1997. Overall, unionization and industrial shifts account for roughly a quarter of the wage losses experienced between 1981 and 1997. This is slightly smaller than implied effects from the decomposition exercise carried out for US wages by DiNardo et al. (1996). Interestingly, none of the wage gains after 1997 can be accounted for by these factors. Combined with Figure 5, these results indicate that compositional shifts are not the main forces driving the decline and then rise in wages that interest us.

Another potential compositional factor is the increasing education level of the workforce in this period. If the "most able" high school educated workers in earlier birth cohorts would have obtained a post-secondary education (and moved out of our sample) in later cohorts, this could explain some of the general wage decline. However, our earlier investigation indicated that the wage patterns do not fit well with birth cohort based explanations. Moreover, such an education selection story cannot explain the increasing profiles after 1997.

4.3.2 Econometric Results

We next turn to presenting the results from our second stage estimation. As discussed earlier, the *Ratchet* and *Below* variables are key covariates in our investigation. Note that we do not observe start wages for all years. However, the gaps in our data (early 1980s, early 1990s), take place during recessions. Given both cyclical considerations and the longer term negative trend that we observe, it seems reasonable to assume that start wages are falling during these periods. As such, the value of *Ratchet* for any cohort in our data will not be determined by start wages in these missing years of data.

Results for various specifications of the second-stage regression are reported in Table 3. In the first column, results are reported for a specification without the *Ratchet* and *Below* variables. The estimated spline coefficients indicate that the profiles essentially become flat after 10 years. The cohort-tenure interactions are consistent with previous plots and with the implicit contracting model outlined in section 2. That model predicts that cohort specific wage profiles will shift down in parallel during the long decline and then shift up and steepen after 1996. The estimated cohort-slope interactions in the first column indicate that profiles are parallel for cohorts entering before the mid–90s[11] with the possible exception of the 1984 cohort, which may have been able to re–negotiate its wages in the late–80s boom. The profiles become much steeper for the later cohorts and the hypothesis that all cohorts share a common slope and spline is strongly rejected ($F_{24,123} = 4.34$, p-value=0.00).

In Column 2, we include the *Ratchet* variable. Its estimated coefficient is 1.13, which implies that a 1% increase in the start wage above the highest start wage experienced so far during a particular job spell results in a 1.13% increase in the wage of that ongoing job spell. The estimated coefficient is not statistically different from one at conventional

[11]The hypothesis that the tenure-wage profiles have the same shape for the 1981–1995 start cohorts cannot be rejected at the 5% level ($F_{14,123} = 1.57$, p-value=0.10).

Table 3: Results, High School Men, Second-stage Regression

Variable	Specification (1)	(2)	(3)
Ratchet		1.128 (0.363)***	0.906 (0.397)**
Below			0.300 (.221)
Tenure	0.013 (0.004)***	0.013 (0.004)***	0.016 (0.004)***
Spline	-0.018 (0.008)**	-0.018 (0.007)**	-0.019 (0.007)**
Cohort Slope Interactions			
1984	0.006 (0.006)	-0.002 (0.007)	-0.002 (0.007)
1987	-0.002 (0.005)	-0.002 (0.005)	-0.002 (0.005)
1989	-0.003 (0.005)	-0.003 (0.005)	-0.003 (0.005)
1995	0.007 (0.005)	-0.003 (0.006)	-0.004 (0.006)
1997	0.011 (0.005)**	0.001 (0.006)	0.000 (0.006)
1999	0.015 (0.006)**	0.006 (0.006)	0.005 (0.006)
2001	0.022 (0.007)***	0.012 (0.008)	0.010 (0.008)
2003	0.040 (0.010)***	0.014 (0.013)	0.016 (0.013)
2005	0.039 (0.020)*	0.021 (0.020)	0.021 (0.020)
Cohort Spline Interactions			
1984	-0.000 (0.013)	0.013 (0.013)	0.011 (0.013)
1987	0.010 (0.010)	0.010 (0.010)	0.005 (0.011)
1989	0.014 (0.010)	0.014 (0.010)	0.009 (0.011)
1995	0.021 (0.028)	0.015 (0.027)	0.014 (0.027)
Cohort Intercepts			
1984	-0.091 (0.029)***	-0.101 (0.028)***	-0.106 (0.028)***
1987	-0.050 (0.027)*	-0.050 (0.026)*	-0.055 (0.026)**
1989	-0.064 (0.029)**	-0.064 (0.028)**	-0.067 (0.028)**
1995	-0.141 (0.030)***	-0.154 (0.030)***	-0.154 (0.030)***
1997	-0.159 (0.027)***	-0.161 (0.026)***	-0.163 (0.026)***
1999	-0.158 (0.028)***	-0.160 (0.027)***	-0.160 (0.027)***
2001	-0.160 (0.029)***	-0.150 (0.028)***	-0.150 (0.028)***
2003	-0.180 (0.031)***	-0.178 (0.030)***	-0.181 (0.029)***
2005	-0.124 (0.032)***	-0.124 (0.031)***	-0.126 (0.031)***
2007	-0.104 (0.033)***	-0.104 (0.032)***	-0.106 (0.032)***
Constant	2.672 (0.022)***	2.672 (0.021)***	2.674 (0.021)***
R^2	0.831	0.842	0.843
N	168	168	168

Notes: *, **, *** indicate different from zero at the 10%, 5%, and 1% levels of significance. Standard errors are reported in parentheses. With the exception of 1984, intercepts and slope/spline interactions for even years are not reported here. See the web appendix for even year estimates. Estimated using WLS, with the inverse of the standard error from the first stage regressions used as weights. First stage regressors consist of 4 age indicators, 13 industry indicators, 9 province indicators and 168 cohort-tenure indicators. The cohort-tenure indicators are the depend variable for the above regressions.

significance levels. Inclusion of the *Ratchet* variable reduces the slope coefficients of the post-97 entry cohorts by a third or more, and it is no longer possible to reject the hypothesis that the underlying shape of the wage-tenure profiles are the same across all job cohorts ($F_{23,122}$=1.12, p-value=0.34). These results suggest that re-contracting accounts for much of the apparent steepening of the wage profiles for the most recent entry cohorts as well as for the 1984 cohort.

In Column 3, we include the *Below* variable as a regressor. This variable captures contemporaneous movements in the entry wage that are below the maximum experienced during an ongoing job spell. In section 2, we noted that an implication of the spot market is that wages in ongoing spells should be as responsive to this variable as they are to the ratchet variable. With the inclusion of this variable, the *Ratchet* variable continues to be statistically significant, and although somewhat smaller in magnitude than the estimate obtained in column 2, the estimated coefficient is still not significantly different from one. In contrast, the coefficient on *Below* is not significant, and has a magnitude close to zero. Once more, the hypothesis that all cohorts share a common wage-tenure profile cannot be rejected ($F_{24,121}$=.80, p-value=0.75). These results are consistent with the re-contracting model.[12][13]

As we noted in Section 2, this specification is consistent with a model of general human capital. If human capital is firm-specific, then recontracting should only occur when the entry wage exceeds the wage in an ongoing spell, inclusive of returns to job tenure. The difficulty in implementing this, of course, is that we cannot determine when re-contracting is needed by comparing entry wages to the actual wage at tenure, j, for a given cohort since the latter wage already reflects any re-contracting. Instead, we need to construct a wage using a tenure profile unaffected by renegotiation. To do this, we used data for the 1981 cohort alone. Of all the cohorts we observe, this is the one least likely to experience recontracting, as it enters the job market under the most favorable conditions of all the

[12]To check further on age composition effects, we also ran a single-stage regression with the controls used in the first stage, along with cohort-specific wage-tenure profiles, full interactions between age groups and cohort intercepts, and full interactions between age groups and tenure. The results are consistent with what we have reported so far, with steepening profiles for more recent cohorts, and the same pattern of decline and partial recovery in entry wages. Results are available in the online appendix.

[13]We obtain broadly similar results working with women, which are reported in the online appendix. Wage profiles steepen for the post-97 cohorts and the coefficient on the Ratchet variable is positive and significant.

cohorts we use. If the recontracting model holds, this cohort can thus be used to obtain an estimate of the "pure" wage-tenure profile.

Estimating a wage equation including only the 1981 entry cohort results in a tenure coefficient of .017 (with a standard error of .002) and a spline interaction after the first ten years of -.023 (with standard error of .005).[14] We also estimated entry wages for each cohort using the same set of controls, but using only workers with up to one year of job tenure and a set of start year dummies. We used these two sets of estimates to construct a set of counterfactual wage-tenure profiles for each cohort. These wages were then compared with the entry wages of subsequent entry cohorts. For the 149 observations exceeding one year of job tenure, there were only 9 cases where the constructed wage profiles did not exceed a subsequent entry wage. This implies that if implicit contracts are to account for the general steepening of the wage profiles after 1997, then the returns to job tenure must represent general human capital or matching since a specific human capital framework does not imply adequate opportunities for recontracting to account for the observed steepening. To emphasize this point, note that entry wages increased by .13 log points between 1995 and 2007, or approximately .010 log points a year. As the return to a year of job tenure was .017 log points a year for the 1981 cohort, wage growth at that rate would result in wages in ongoing spells easily outpacing the changes in entry wages for subsequent cohorts after 1997.

4.4 Robustness of the Results

The results reported above were obtained using cohort-specific spline functions for the wage-tenure profile in the second-stage regression. We next investigate the sensitivity of our results to alternative specifications of the wage-tenure profile. In the existing literature evaluating the role of implicit contracts to account for wages over the business cycle, a wide array of functional forms are used. Beaudry and Dinardo (1991) use a linear function, Grant (2003) uses a quadratic function, while Devereux and Hart (2007) use a cubic function. To examine the robustness of our results to alternative specifications, we estimated a model

[14]The regression includes a spline specification for the tenure profile and controls for union status, industry, age, and province of residence.

which imposes a common non-parametric wage-tenure profile for all cohorts. In a specification including the *Ratchet* variable but not *Below*, the estimated coefficient on *Ratchet* was 1.30 (s.e. of .15); while in a specification including both *Ratchet* and *Below*, the estimated coefficients were .95 (s.e. of .18) and .43 (s.e. of .12), respectively.[15]

The *Ratchet* and *Below* coefficients from the common tenure profile specification are not strictly consistent with any of the three models outlined in Section 2. However, they do imply that the wages in ongoing spells only partially track conditions in the labour market, with increases in start wages above the previous maximum in the job spell having a much stronger effect on wages in that spell than other movements in the current start wage. These results mirror the findings in the business cycle literature. Only Beaudry and Dinardo (1991) find evidence that the tightest labour market conditions during a job spell fully determines the current wages. Both Grant (2003) and Devereux and Hart (2007) find a significant and negative coefficient on the contemporaneous unemployment rate as well. McDonald and Worswick (1999) obtain similar results for Canadian nonimmigrant males, but they also find a positive coefficient on the unemployment rate at the time of entry; this result may reflect the non-stationarity of Canadian wages during their study period (1980–92). Although we are examining non-cyclical wage patterns, our results also indicate that both the tightest labour market during a job spell and the current labour market play roles in the behaviour of the wages of high school men. Given the long term nature of the decline, it may be that firms faced bankruptcy constraints that prevented them from fully insuring against such a decline. Gamber (1988) uses a two-period model to illustrate that such constraints may result in wage contracts that respond asymmetrically to permanent and temporary shocks. In particular, in states where the bankruptcy constraint is binding, firms are unable to insure fully.

[15] A visual examination of the non-parametric specification suggests that a cubic profile describes the data well.

5 The Role of Unionization Revisited

So far, our results are based on a simple regression framework in which union membership results in a common premium for all workers.[16] Empirical studies on union wage effects indicate the returns to experience tend to be lower in the union sector (Kuhn, 1998). If the union premium is higher for entrants, and if non-union workers experience a higher return to tenure, then the falling entry wages and steepening wage-tenure profiles we observe may be driven by effects of de-unionization that are not taken into account in this model, rather than by recontracting. To examine this issue further, we estimate our model separately for union and non-union workers. We begin with cohort-specific spline functions for the wage-tenure profile. Results are presented in Table 4.

The results for the non-union workers are similar to those for the pooled sample. In the absence of the *Ratchet* variable (Column 1), the wage profiles steepen for the most recent entry cohort. The cohort intercepts show the same pattern of decline until the 1990s, followed by signs of partial recovery in wages. The wage-tenure profiles are steeper than those in the pooled sample, and while they become flatter after 10 years of job tenure, the point estimate implies further wage growth at longer lengths of job tenure. Inclusion of the variables of interest (Column 2) results in a somewhat larger estimate of the Ratchet coefficient, but little indication that wages respond to outside conditions, except to recontract to match a wage that exceeds the highest during the job spell. Again, once the implicit contract variables are included, it is not possible to reject the hypothesis that the wage profiles have identical slopes across cohorts ($F_{24,121}$=.90, p-value=0.60). The point estimate of the *Ratchet* coefficient implies overbidding to retain current employees, but again the estimate is not significantly different from one.

Following the procedure that was used for the pooled sample, we estimated our model for the non-union sector imposing a common, non-parametric specification for the wage-tenure profile. As with the pooled sample, this profile could be closely approximated with a cubic. The estimates of the coefficients of Ratchet and Below were .95 and .43 respectively, with standards errors of .24 and .23. These estimates are similar to the estimates that we

[16] The estimated union premium is .171.

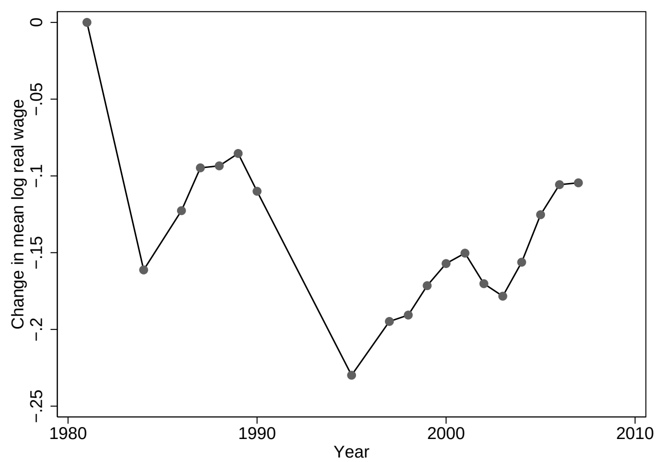
Table 4: Results, High School Men, Second-stage Regression, Union and Non-Union

	Specification			
	Non-Union		Union	
Variable	(1)	(2)	(3)	(4)
Ratchet		1.461 (0.497)***		0.071 (0.244)
Below		0.119 (.309)		0.123 (0.107)
Tenure	0.017 (0.005)***	0.019 (0.006)***	0.010 (0.004)**	0.011 (0.004)**
Spline	-0.007 (0.010)	-0.008 (0.010)	-0.024 (0.008)***	-0.022 (0.008)***
Cohort Slope Interactions				
1984	0.010 (0.008)	-0.006 (0.009)	-0.008 (0.007)	-0.007 (0.007)
1987	-0.004 (0.006)	-0.005 (0.006)	-0.006 (0.005)	-0.005 (0.005)
1989	-0.001 (0.006)	-0.002 (0.006)	-0.011 (0.005)**	-0.009 (0.005)*
1995	0.007 (0.007)	-0.009 (0.008)	-0.003 (0.006)	-0.003 (0.006)
1997	0.012 (0.006)*	-0.002 (0.008)	0.001 (0.006)	-0.000 (0.006)
1999	0.017 (0.007)**	0.004 (0.008)	0.008 (0.006)	0.007 (0.007)
2001	0.019 (0.009)**	0.005 (0.010)	0.019 (0.009)**	0.016 (0.009)*
2003	0.043 (0.013)***	0.011 (0.016)	0.020 (0.013)	0.014 (0.014)
2005	0.049 (0.025)*	0.031 (0.024)	-0.014 (0.026)	-0.023 (0.027)
Cohort Spline Interactions				
1984	-0.020 (0.017)	0.007 (0.019)	0.028 (0.014)**	0.026 (0.014)*
1987	0.007 (0.014)	0.008 (0.014)	0.015 (0.011)	0.012 (0.011)
1989	0.001 (0.014)	-0.000 (0.014)	0.030 (0.011)***	0.026 (0.011)**
1995	0.009 (0.036)	0.008 (0.035)	0.049 (0.030)	0.038 (0.031)
Cohort Intercepts				
1984	-0.126 (0.037)***	-0.148 (0.037)***	0.008 (0.033)	0.007 (0.032)
1987	-0.062 (0.035)*	-0.068 (0.034)**	-0.007 (0.030)	-0.008 (0.030)
1989	-0.086 (0.037)**	-0.086 (0.035)**	-0.002 (0.032)	-0.006 (0.032)
1995	-0.148 (0.038)***	-0.167 (0.037)***	-0.082 (0.036)**	-0.087 (0.036)**
1997	-0.160 (0.034)***	-0.163 (0.033)***	-0.133 (0.031)***	-0.134 (0.031)***
1999	-0.152 (0.035)***	-0.154 (0.034)***	-0.164 (0.033)***	-0.166 (0.033)***
2001	-0.147 (0.036)***	-0.135 (0.035)***	-0.167 (0.035)***	-0.160 (0.035)***
2003	-0.169 (0.038)***	-0.175 (0.037)***	-0.202 (0.037)***	-0.197 (0.037)***
2005	-0.119 (0.040)***	-0.123 (0.039)***	-0.117 (0.039)***	-0.115 (0.039)***
2007	-0.107 (0.041)**	-0.108 (0.040)***	-0.033 (0.041)	-0.034 (0.041)
Constant	2.666 (0.028)***	2.667 (0.027)***	2.835 (0.024)***	2.835 (0.024)***
R^2	0.830	0.842	0.709	0.711
N	168	168	168	168

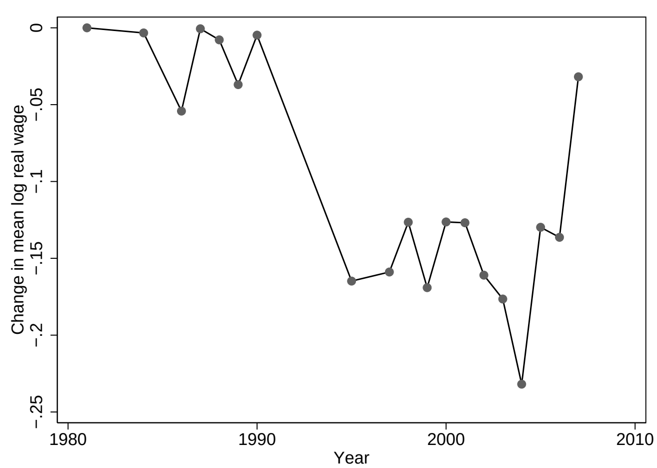
Notes: *, **, *** indicate different from zero at the 10%, 5%, and 1% levels of significance. Standard errors are reported in parentheses. With the exception of 1984, intercepts and slope/spline interactions are not reported for even years. Estimated using WLS, with the inverse of the standard error from the first stage regressions used as weights. First stage regressors consist of 4 age indicators, 13 industry indicators, 9 province indicators and the 168 entry-tenure indicators used as the dependent variable in the second-stage.

obtained using the pooled sample, and again indicate asymmetries in how wages in ongoing spells respond to various changes in labour market conditions. Wages in ongoing spells adjust fully when entry wages rise above the highest wage experienced so far during the spell, but only decrease partially when wages for entrants fall.

Figure 6: Entry Wages for High School Men, Relative to 1981, By Union Status



(a) Non-union



(b) Union

The results for unionized workers imply a much different wage pattern over time. When the implicit contract variables are not included (column 3), we see that union entry wages did not start falling until after 1990. Wage-tenure profiles are also flatter and, unlike the non-union sector, there is little evidence that wage profiles steepened for the most recent cohorts. Consistent with these finding, the coefficients on the contracting variables are small and not significantly different from zero (Column 4).

The estimated contract variable effects for union workers would seem to be consistent with the full commitment model. However, the *Ratchet* and *Below* variables used to obtain the union results were computed using entry wages in the union sector, which followed a somewhat different pattern that those for the non-union sector. In Figure 6, we plot composition-controlled entry wages for union and non-union workers. As expected, given the predominance of non-union jobs at entry, entry wages for non-union workers closely resemble the entry wage for the pooled sample. For both the non-union and pooled samples, variation in *Ratchet* comes from the recovery from the early 1980s recession and from the general upswing after 1995, while variation in *Below* comes from the same recession, along with the decline between 1990 and 1995, and the economic slow-down in the early 2000s. In the union sector there is little variation in entry wages with which to identify the *Ratchet* coefficient, and the only substantive variation in *Below* comes from the decline in entry wages between 1990 and 1995. Thus, our results for the union sector appear to be driven by the lack of useful variation in entry wages over our study period and it is difficult to support specific conclusions from them.

6 Technological Change and Entry Wage Movements

To this point, we have established that an implicit contract model with renegotiation provides a reasonably good fit for the wage patterns of Canadian less skilled workers. This in turn implies that understanding wage movements in this market comes down to understanding what is driving patterns in entry wages. Our earlier decomposition results indicate that shifts in unionization and industrial composition played only a small role in determining the entry wage movements – accounting for only a quarter of the decline between 1981 and 1997, and none of the post-1997 recovery. In this section, we briefly examine whether technological change can explain the remaining movements in entry wages.

Research over the last 15 years on technological change and the labour market suggests that a combination of relative factor supply movements and technical change could help understand wage movements beyond those due to compositional shifts (e.g., Katz and Murphy (1992) and the large volume of research that followed it). Thus, consider a constant returns to scale aggregate production function with inputs being unskilled labour, U, skilled labour, S, and physical capital, K. If we allow for technical change that enhances unskilled and skilled labour at potentially different rates, then equating the marginal product for unskilled workers to their wage and taking a log-linear approximation yields a simple equation:

$$ln(w_t^U) = \alpha_0 + \alpha_1 ln(\frac{S_t}{K_t}) + \alpha_2 ln(\frac{U_t}{K_t}) + \alpha_3 t + e_t \tag{14}$$

where α_3 reflects the impact of skill biased technical change on unskilled wages.

In the first column of Table 5, we present estimates of this equation where we use the average wage of nonunion, high school or less educated workers with one year or less of job tenure for the relevant unskilled wage and use measures of aggregate unskilled and skilled labour and physical capital described in Appendix B. We use the nonunion entry wage to control for shifts in union composition.[17] We use annual observations on each of the variables and so have only 19 observations. While this is a small number, our results from the previous sections indicate that understanding the low skilled wage structure comes down to understanding these observations. Our estimates are obtained employing weights corresponding to the number of observations underlying the average wage in each year.

The resulting estimates indicate that an increase in unskilled labour leads to an increase in the unskilled wage. This is the opposite of what theory predicts but the estimated coefficient is close to zero and statistically insignificant. The coefficient on the time trend is also small and statistically insignificant at any conventional significance level. In contrast, the coefficient on the ratio of skilled labour to physical capital has a strong and statistically significant negative effect. The R^2 for the regression is .57 and only falls to .56 when we remove the time trend. Thus, about 60% of the variation in the nonunion entry wage is accounted for by relative factor movements.

The wage regression reported in column one has the potential for endogeneity problems because individuals choose their education based on expected wages associated with different

[17] Results when we use union starting wages are similar in spirit, again showing a significant negative effect from $\frac{S_t}{K_t}$, an insignificant negative effect from $\frac{U_t}{K_t}$ and a small and statistically insignificant time effect.

Table 5: The Impact of Relative Factor Movements on Nonunion Entry Wages

	No TFP Correction		With TFP Correction	
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
Constant	-14.59 (13.01)	-18.83 (32.35)	-15.00 (7.23)*	-31.82 (12.69)**
$\ln(S_t/K_t)$	-.77 (.22)***	-.94 (.34)**	-.62 (.30)*	-1.21 (.45)**
$\ln(U_t/K_t)$	.033 (.61)	.022 (1.69)	-.14 (.25)	-.27 (.38)
Trend	.014 (.014)	.017 (.039)	.0043 (.011)	.022 (.015)
Number of Obs	19	19	19	19
R^2	.57	-	.91	-

Notes: *, **, *** indicate different from zero at the 10%, 5%, and 1% levels of significance. Standard errors are reported in parentheses. All specifications employ weights equal to the sample size underlying the calculations of the average nonunion wage.

education levels and investment decisions are based on factor price patterns, implying that the factor ratio variables may be correlated with the error term. Beaudry and Green (2005) argue that shifts in the age distribution of the population are plausible instruments for the factor ratios because the movement of the baby boom through the age structure (and its replacement of earlier, less educated cohorts) will shift the stocks of U and S but, unless fertility decisions in the years just after WWII were based on accurate predictions of wage movements decades later, they are not expected to be related to current wage movements. Further, given that investment and savings vary with age, shifts in the age structure may also determine capital formation patterns. The estimates presented in columns 2 and 4 use the proportion of the population in ten year age groups from 20 to 60 as well as the proportion aged 60 and over as instruments for S_t/K_t and U_t/K_t.[18] The column 2 results indicate that addressing endogeneity in this way does not change the basic conclusions from column 1: the effect of S_t/K_t is large, negative and statistically significant, while the effects of both U_t/K_t and time are small and statistically insignificant.

The specific values taken by the estimated coefficients reflect a similar pattern to those obtained by Beaudry and Green (2005) for the US. In their derivation of the low skilled

[18]The instruments perform well in the first stage. The F statistic associated with the joint hypothesis that the age proportion variable effects are zero is 129.94 for the S_t/K_t regression. The same statistic for the S_t/K_t regression is 35.75. Given that the 99th percentile critical value for these statistics (which are distributed F(4,11)) is 5.67, the hypothesis that the instruments have a zero effect in the first stage is easily rejected.

wage equation, they allow for possible effects of general productivity movements captured in TFP measures on relative factor supplies.[19] When we use their adjusted variables, we obtain the results presented in column 3. These are quite similar to their results for the US, where the coefficients (standard errors) they estimate in their low skilled wage equation are -.69 (.31) and -.0020 (.32) for their S/K and U/K variables, respectively, and .0059 (.012) on the time trend. They argue that these coefficients fit with an induced technical change model in which firms can choose between a skilled and an unskilled intensive technology.[20] In that case, the coefficients imply that increases in the supply of skilled labour induce firms to choose the skill intensive technology. This, in turn, leads to less capital being applied to unskilled workers and declines in their wages. Thus, the key variable in determining wage movements is the relative supplies of skilled labour and capital, as we see in our estimates. Finally, in column 4, we use the TFP adjusted variables but instrument using the same instruments as in column 2. Again, the instruments perform well in the first stage and the results are similar in spirit to those from OLS. The coefficients estimated using instrumental variables are all larger in magnitude than those obtained from OLS.

The results presented in Table 5 do not fit with some other notable models of wage impacts of technical change. In particular, the model set out in Katz and Murphy (1992), which involves a more restrictive specification for the aggregate technology, implies that holding U and K constant, an increase in S should lead to an increase in the unskilled wage as unskilled labour becomes relatively scarce. This effect corresponds to α_1 in the above specification. Our estimates of α_1 are strong and negative: the opposite of the Katz-Murphy prediction. Similarly, Krusell et al. (2000) argue that capital-skill complementarity and an increase in the effective capital stock explain wage patterns in the US in this period. However, their model implies that an increase in capital should lead to declines in the unskilled wage, holding U and S constant. In the estimates here, the partial effect of capital corresponds to $-(\alpha_1+\alpha_2)$, which has a strong positive sign: the opposite of the results that Krusell et al. (2000) obtain with a more restrictive production specification. Indeed, while

[19] Specifically, they use adjusted ratios, $ln(S_t^*/K_t) = ln(S_t/K_t) + TFP_t/(s_t^U + s_t^S)$ and $ln(U_t^*/K_t) = ln(U_t/K_t) + TFP_t/(s_t^U + s_t^S)$, where s^U and s^S are expenditure shares for skilled and unskilled labour and TFP is a consistent total factor productivity measure. They also adjust the unskilled wage as $ln(w_t^U *) = ln(w_t^U) - TFP_t/(s_t^U + s_t^S)$.

[20] This model is similar in spirit to those in Basu and Weil (1998) and Caselli (1999).

the arguments in these earlier papers imply that increases in education will have beneficial effects for the less skilled because they become a relatively scarce factor, the estimated results here (and the model in Beaudry and Green (2005)) point to the opposite conclusion. In the technology selection model, as long as the economy remains in the period of transition in which firms are choosing between the two types of technologies, increasing the education level will benefit those who obtain the education but the remaining unskilled will have less and less capital to work with and will be left further behind as the number of skilled workers increases. What will help unskilled workers, given these estimates, is increases in physical capital. This, broadly speaking, would imply that the expanding group of skilled workers can be outfitted without taking capital away from the unskilled.

Given the predominant role of the S/K variable in the estimates in Table 5, the implication is that we can understand movements in entry wages in terms of productivity movements induced by relative movements in human and physical capital. During the long decline, education levels were rising faster than capital per worker but after 1997 this was reversed as capital flowed into Canada and expansion in the stock of skilled labour slowed.

In this section, the model underlying equation (14) assumes that wages represent the marginal product of unskilled labour. If wages are indeed governed by implicit contracts, as we have argued in the previous sections, entry wages in any given year will depend on a combination of productivity in that year and expectations about future productivity, meaning that entry wages will not be perfectly correlated with current period marginal product. Thus, we view these results as suggestive; potentially fitting with a notion that firms and workers are relatively myopic in productivity projections.

7 Conclusion

In this paper, we explore the wage patterns of high school educated males in Canada. The last three decades have witnessed a dramatic drop in their real wages up to the mid–1990s followed by a recovery that has still left them with real wages well below their 1981 peak. The time pattern was quite different for workers with different amounts of job tenure, however. Workers whose jobs started in the early 1980s, and hence had high job tenure in the 1990s and 2000s, had wages that were seemingly unaffected by downturns and upturns in the wages of new job starters. Our primary interest is in understanding this pattern. We argue that the wage-tenure patterns for less skilled Canadian workers over the period 1981–2007 fits well with an implicit contracting model with renegotiation. In such a model, workers are at least partially insured against falls in wages but can ratchet up their wages in strong economic times. This raises important concerns for the flexibility of wages in this market in a period when Canada appears headed into an economic downturn.

Within the context of an implicit contracting model with renegotiaton, understanding overall wage movements comes down to understanding movements in entry wages. We show that the composition of new jobs has shifted over the last twenty-five years, with particularly dramatic declines in unionization levels of new job starters. However, a standard decomposition exercise indicates that these shifts explain only about a quarter of the decline in entry wages between 1981 and 1997 and none of the increase after 1997. Instead, we need to look for explanations for why productivity first declined and then increased for this group. In the last section of the paper, we argue that a model in which technological change is induced by relative factor flows provides a potential explanation. In particular, according to this model, unskilled wages declined because physical capital did not grow fast enough to outfit the expanding stock of skilled workers without reducing the amount of capital applied to unskilled workers. This process was reversed after 1997 as capital inflows outstripped increases in the number of skilled workers. If this is true, it implies that help for unskilled workers may lie mainly in the capital market.

Appendix A The Size of Synthetic Cohorts

Table A-1 shows the number of observations available at several lengths of the job spell for all the cohorts in our data set. The number of observations falls dramatically after the first year, and then continues to taper off gradually over time. These figures do not represent actual attrition from job spells, since the design of the samples differs somewhat. For example, in the 1995 SWA, which we use only for the 1995 job entry cohort, there is no imputation for missing values and only half of the available LFS sample is surveyed; as a

Table A-1: Number of observations, High School Men, By Start Year and Duration of Job Spell, Selected Durations

Start Year	Years of Job Tenure					
	0	1	2	5	10	15
1981	2,476			442		
1984	1,886		574	210		151
1986	2,117	1,519	518			187
1987	2,641	927	455		307	167
1988	2,319	1,066	670		287	163
1989	1,882	1,113			276	160
1990	1,736				251	204
1995	521		582	360	227	
1997	1,759	785	571	355	252	
1998	1,650	716	558	337		
1999	1,730	711	607	360		
2000	1,599	768	623	451		
2001	1,571	742	551	331		
2002	1,434	611	518	373		
2003	1,436	675	517			
2004	1,433	717	592			
2005	1,531	917	599			
2006	1,679	877				
2007	1,631					

result, the number of observations is considerably smaller than other years. Blank entries indicate cases where a spell of a particular length either (i) occurs in a year for which data is unavailable (e.g. 1992) or (ii) has not yet occurred (e.g. 2005 job starters with 5 years of job tenure). The key point from the table is that we have sufficient observations to establish wage patterns even at long job durations.

Appendix B Non-Wage Data

In this appendix, we provide information on the non-wage data series used in the paper. We use net capital stock series (residential and non-residential) from Statistics Canada for our physical capital stock measure. For both high and low skilled workers, we construct measures of total hours worked using the SCF from 1981 through 1997. In particular, for each worker, we construct total hours worked by multiplying weeks worked in the previous year times hours worked in the reference week. We then sum these hours (multiplying by the weights provided in order to provide an estimate at population level) for each of three education groups: high school or less; some or completed post-secondary (less than a BA); and BA or higher university education. The education categories were reformed in 1990. To address this, we estimated a simple regression of the hours series on a cubic in time plus a dummy variable equalling one for 1990 and after. We find that this approach - with a common time trend before and after 1990 combined with a step up or down - fit the various data series very well. We subtract the coefficient on the post-1990 dummy variable from the total hours series in all post-1990 years to construct a smoothed series for total hours for each education group. We use SLID data for the years 1997-2005 and the LFS for the years 2006–07 in a similar way to construct the hours series for those years, using the overlap years of 1997 and 2004/05 to normalize the SLID and LFS based series to the SCF based series. We need to combine our three education categories into two skill groups. To do this, we follow Katz and Murphy (1992), creating a low skilled group that consists of the high school or less education category plus .44 times the some post-secondary category. The high skilled group contains the university educated plus .56 times the some post-secondary group.

Finally, we require a TFP series which is consistent with the three factor (skilled labour, unskilled labour and capital) production function we employ. To do this, we construct a weighted average of growth in each of the inputs in each year, using the share of each factor in total income in 1981 as the weights. We then subtract that weighted average from the growth in GDP for each year to generate the annual change in TFP and then cumulate those changes to get a TFP level measure (measured relative to 1981).

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