# AN EXPLORATION IN JOB STABILITY 

by
PIERRE BROCHU
B.A., The University of Toronto, 1989

M:A., The University of Toronto, 1990

# A THESIS SUBMITTED IN PARTIAL FULFILMENT OF THE REQUIREMENTS FOR THE DEGREE OF <br> DOCTOR OF PHILOSOPHY <br> in <br> THE FACULTY OF GRADUATE STUDIES 

(Economics)

THE UNIVERSITY OF BRITISH COLUMBIA
November 2006
(C)Pierre Brochu, 2006

## Abstract

The goal of this dissertation is to document striking new job stability patterns in Canada and explore their causes.

The first paper (Chapter 2) shows how to correctly apply the retention rate approach to cross sectional data. I propose two cross sectional estimators and clearly identify the conditions required for consistency. I demonstrate the bias of existing approaches for calculating standard errors, and propose an alternative method. Finally, using Current Population Survey data I show that existing approaches to estimating standard errors may lead the researcher to falsely reject the null hypothesis of no change in job stability.

The second paper (Chapter 3) documents the changing job stability patterns in Canada over the 1977-2004 period. I use a retention rate approach which is less sensitive to job inflows than in-progress measures, but for which the data requirements are severe. In, North America, only the Canadian Labour Force Survey satisfies these stringent data requirements. Using this rich source of data and tools developed in Chapter 2, I find that overall job stability has actually increased in Canada since the early 1990s. Two other key findings include an increase in the relative stability of women and a large increase in stability for jobs with initial tenure of less than one year.

The third paper (Chapter 4) explores the causes of the new job stability patterns that were documented in Chapter 3. Results indicate that the ageing of the workforce, increased educational attainment and increased labour force attachment of women play an important role in the aggregate patterns. However, only rising educational attainment matters for newer jobs-with a large part of the increase still unexplained. I use a match quality framework to explore for changes in job stability along tenure lines. The model predicts that stricter eligibility requirements introduced in the early 1990s for the Employment Insurance program should result in better match formation, and as such, lead to an increase in stability of newer jobs. The empirical findings of this thesis are consistent with such predictions.

## Table of Contents

Abstract ..... ii
Table of Contents ..... iii
List of Tables ..... v
List of Figures ..... vi
Acknowledgements ..... viii
Dedication ..... ix
1 Overview and Summary ..... 1
2 Retention Rate Approach for Repeated Cross Sectional Data ..... 5
2.1 Introduction ..... 5
2.2 The Literature ..... 6
2.3 Empirical Approach ..... 15
2.4 Population Retention Rate ..... 15
2.5 Proposed Cross Sectional Estimator ..... 17
2.6 Panel Estimator ..... 19
2.7 Comparison of Panel and Cross Sectional Estimator ..... 21
2.8 Application to Survey Data ..... 22
2.9 Existing Cross Sectional Methods and their Biases ..... 24
2.10 Testing for Changes in Job Stability ..... 26
2.10.1. Application to CPS data ..... 30
2.11 'Averaged' Retention Rate ..... 33
2.12 Conclusion ..... 38
3 Job Stability Patterns ..... 39
3.1 Introduction ..... 39
3.2 LFS Data ..... 40
3.3 Canadian Job Stability Patterns ..... 42
3.3.1 Controlling for the Business Cycle ..... 60
3.4 Canada - U.S. Comparison ..... 63
3.5 Conclusion ..... 69
4 Explaining the New Canadian Job Stability Patterns ..... 70
4.1 Introduction ..... 70
4.2 Retention Rate Tools ..... 71
4.2.1 Aggregate Patterns ..... 72
4.2.2 Sub-Population Patterns ..... 74
4.3 Source of the Aggregate Patterns ..... 77
4.4 Sources of the Low Tenured Worker Pattern ..... 86
4.4.1 Compositional Changes ..... 87
4.4.2 Employment Insurance Reform ..... 90
Match Quality Search Model ..... 91
Results ..... 95
4.5 Conclusion ..... 96
5 Concluding Remarks ..... 98
Bibliography ..... 100
Appendices ..... 104
Appendix A: Education Categories ..... 105
Appendix B: Data Related Issues ..... 106
Appendix C: Asymptotic Properties ..... 109

## List of Tables

2.1 LFS Adjustment factor ( $a d j_{j}$ ) by Year ..... 23
2.2 4-year Male Retention Rates - Time Differentials ..... 31
2.3 4-year Female. Retention Rates - Time Differentials ..... 32
2.4 4-year Retention Rates - Group Differentials ..... 34
3.1 Comparison Across Business Cycles ..... 61
3.2 Comparison Across Business Cycles, by Sector and Region ..... 62
3.3 Comparison Across Business Cycles - Conditioning on Initial Tenure Less than 1 Year ..... 63
3.4 Canada-U.S. Differential in the year 2000 ..... 68
4.1 Decomposition Across Business Cycles: 1987-88 and 1998-99 ..... 81
4.2 Decomposition Across Business Cycles: 1987-88 and 1998-99 ..... 85
4.3 Decomposition Across Business Cycles: 1987-88 and 1998-99 ..... 86

## List of Figures

3.1 One Year Retention Rate ..... 43
3.2 One Year Retention Rate, by Gender ..... 44
3.3 One Year Retention Rate, Male-Female Differential ..... 45
3.4 One Year Retention Rate, by Age ..... 46
3.5 One Year Retention Rate, by Education ..... 47
3.6 One Year Retention Rate, by Industry ..... 48
3.7 One Year Retention Rate, by Industry Breakdown ..... 49
3.8 One Year Retention Rate, Public versus Private ..... 49
3.9 One Year Retention Rate, by Region (West) ..... 50
3.10 One Year Retention Rate, by Region (East) ..... 50
3.11 One Year Retention Rate, by Region (East) ..... 51
3.12 One Year Retention Rate, Less than One Year of Tenure ..... 52
3.13 One Year Retention Rate, One or More Years of Tenure ..... 53
3.14 One Year Retention Rate, One or More Years of Tenure breakdown ..... 53
3.15 One Year Retention Rate, Low Initial Tenure ..... 54
3.16 One Year Retention Rate, Less than One Year of Tenure - by Gender ..... 55
3.17 One Year Retention Rate, One or More Years of Tenure - by Gender ..... 56
3.18 One Year Retention Rate, Less than One Year of Tenure - by Age ..... 57
3.19 One Year Retention Rate, One or More Years of Tenure - by Education ..... 57
3.20 One Year Retention Rate, Less than One Year of tenure - by Industry ..... 58
3.21 One Year Retention Rate, Less than One Year of Tenure - by Industry Breakdown ..... 58
3.22 One Year Retention Rate, Less than One Year of tenure - Public versus Private ..... 59
3.23 One Year Retention Rate, Less than One Year of Tenure - by Region (West) ..... 59
3.24 One Year Retention Rate, Less than One Year of Tenure - by Region (East) ..... 60
3.25 Four Year Retention Rate, Canada - U.S. Comparison ..... 66
3.26 Four Year Retention Rate, Canada - U.S. Comparison - Males ..... 66
3.27 Four Year Retention Rate, Canada - U.S. Comparison - Females ..... 67
3.28 Four Year Retention Rate, Canada - U.S. Comparison - Three Years or Less of Initial Tenure ..... 67
4.1 Industry Decomposition ..... 78
4.2 Education Decomposition - Overall ..... 79
4.3 Education Decomposition - Male ..... 80
4.4 Education Decomposition - Female ..... 80
4.5 Age and Gender Decomposition ..... 83
4.6 Age Decomposition - Males ..... 84
4.7 Age Decomposition - Females ..... 84
4.8 Age and Education Decomposition - Males ..... 86
4.9 Age Decomposition - Males and Less Than one Year of Initial Tenure ..... 88
4.10 Age Decomposition - Males and One or more Years of Initial Tenure ..... 88
4.11 Education Decomposition - Males and Less Than One Year of Initial Tenure ..... 89

## Acknowledgements

I would like to thank David Green, my thesis supervisor, for his guidance and encouragement throughout the thesis process. David patiently answered my many questions over the years for which I am very grateful. I am also indebted to Patrick Francois, Thomas Lemieux and Kevin Milligan, my committee members, for their ongoing support. I appreciated their willingness to meet with me at short notice.

I thank TARGET for providing me with the financial means to complete this thesis.
I thank my fellow graduate students, Aneta Bonikowska, Doris Poon and Tammy Schirle. I enjoyed our discussions and welcomed the opportunities to broaden my understanding. Finally, many thanks to Joris Pinkse whose enthusiasm for econometrics proved to be contagious.

## Dedication

This thesis is dedicated to my wife Karen whose support and belief in me made this dream possible.

## Chapter 1

## Overview and Summary

It is widely acknowledged that technological change, combined with an increase in international trade, has had a significant impact on the economy. The new economy literature has emphasized how these forces have altered the employer-employee relationship, resulting in a breakdown of traditional job arrangements, and a rise in non-standard work, i.e. temporary work and self-employment (e.g., Autor (2003); Vosko, Zukewich, and Cranford (2003)). A perception exists that, in this new economy, workers are now disposable just like any other resource-implying that job stability has declined. However, researchers who directly examined this issue found no evidence of a long term decline in job stability. This mainly U.S. based research faces significant data limitations, and in many cases, the tools used to estimate job stability lack precision. These shortfalls make it difficult to differentiate between cyclical and secular change. Using more refined econometric tools, and a rich source of Canadian tenure data, I find that job stability has actually increased. The objective of this thesis is to document the striking new Canadian job stability patterns and explore their causes.

Examining job stability will facilitate a better understanding of the work relationship. For the majority of families in Canada, labour is their main resource, and generates most of their income. For individuals that work, a large fraction of their waking hours is spent in the workplace. Therefore, a better understanding of the work relationship will lead to a better understanding of what matters to households. Changes in job stability will have public
policy implications. Given that pensions are typically associated with long term employeeemployer relationships, changes in job stability will affect the efficient design of retirement income systems, e.g. RRSPs. The appropriate mix of active labour market programs will depend on the source of the changes, whether due to demographics or behavioral changes of workers. Finally, in a world of incomplete contracts, a decrease in job stability may lead to lower productivity and thus be detrimental to economic growth (e.g. Francois and Roberts (2003) and Ramey and Watson (1997)).

In Chapter 2, I present the empirical approach used in this thesis. The approach is based on the retention rate, i.e. the probability that a job of a particular length will continue one more period. The retention rate approach has two important advantages. One, there is a direct link between job stability and retention rates. A job with the same employer is less stable if it has a lower probability of lasting one more period, i.e. if the retention rate falls. Two, retention rates that condition appropriately have the advantage of being less sensitive to job inflows than in-progress job spell measures. ${ }^{1}$ Job inflows are an important issue for Canada, which experienced a historically large increase in labour force participation of women and a significant demographic change brought about by the baby-boomers.

This chapter shows how to correctly apply the retention rate approach to cross sectional data. I show that one can represent the population retention rate as a ratio of two population means-a representation that is easily applicable to cross sectional data. I propose two cross sectional estimators and provide a discussion of their respective identifying assumptions and asymptotic properties. The cross sectional literature (e.g. Neumark, Polsky, and Hansen (2000) and Swinnerton and Wial (1996)) uses methods developed for panel data sets. These methods assume that individuals can be followed over time--an untenable identifying assumption for repeated cross sections. The construction of standard errors is particularly problematic. Existing variance estimators do not fully account for the variability resulting from the inability to follow individuals over time in repeated cross sections, and I show that this will lead to a downward bias. Finally, using Current Population data, I show that the bias of existing approaches can lead the researcher to falsely reject the null

[^0]hypothesis of no change in stability.
In Chapter 3, I document job stability patterns in Canada over the 1977-2004 period. This thesis uses the master Labour Force Survey (LFS) files for the 1977-2004 period as repeated cross sections. In order to examine long term changes in job stability using retention rates, detailed and consistent tenure data available on a regular basis are essential. The Canadian LFS is the only North American data set that satisfies the stringent data requirements of the retention approach. A consistent job tenure question has been part of the monthly LFS starting in 1976. As a result, a full set of one year retention rates can be constructed from the late 1970s through to the mid 2000s.

Using this rich data set and tools developed in Chapter 2, I find that what was previously seen as cyclical change is actually a secular increase in job stability. Since the early 1990s, job stability increased at the aggregate level, and especially for women and workers with low tenure jobs. For the latter group, the change is particularly striking. Comparing the 19871989 and 1988-2000 periods, both strong expansionary periods, job stability for workers with less than one year of initial tenure increased from $44.4 \%$ to $54.5 \%$.

In Chapter 4, I explore the causes of the new job stability patterns identified in Chapter 2. The detailed accounting includes the decline of unionization, ageing of the workforce, increased labour force attachment of women, changing industry structure, and employment insurance reform.

To quantify the importance of compositional changes on the overall retention rate, I use a standard Oaxaca-Blinder decomposition approach. Results indicate that ageing of the workforce, increased educational attainment and increased labour force attachment of women play an important role in the new aggregate job stability patterns that emerge. In particular, the first can explain approximately half of the increase in overall stability from the late 1980s to the late 1990s. This Chapter offers an extension to the Oaxaca-Blinder approach. This extension measures the effect of compositional changes in the workforce on more narrowly defined retention rates. Using this method, I find that ageing of the workforce cannot explain the large increase in stability of newer jobs; the only compositional change that matters is education.

I explore the dramatic changes in stability along tenure lines through the lens of a match quality framework, where a job is both an inspection and experience good. Within such a framework, the stricter EI eligibility requirements introduced in the early 1990s should : make job seekers more selective in their acceptance of job offers, and lead to fewer low quality matches being formed. This is consistent with the findings that the changes are localized at very low levels of initial tenure, and why the increases in stability are more modest in scale once one conditions on having worked more than a few months.

## Chapter 2

## Retention Rate Approach for

## Repeated Cross Sectional Data

### 2.1 Introduction

The job stability literature, which focusses on the probability of a job ending, is part of a larger body of literature that examines labour market transitions. Given a panel data set with job tenure information, one could apply longitudinal tools developed for the exploration of unemployment-employment transitions to identify long term breaks in job stability. For both segments of this broad literature, the object of interest is a transition probability. The examination of long term changes in job stability requires detailed and consistent tenure data over an extended period of time-requirements that are not met in North American data sets. ${ }^{1}$ Job stability researchers who have turned to repeated cross sectional data have nevertheless relied on such longitudinal tools. Yet, these (panel) tools are designed to take advantage of the fact that one can follow individuals over time - something not possible with cross sectional data. As a result, existing cross sectional approaches have had to impose very strict identifying assumptions, and in some cases, make unverifiable claims that their

[^1]approaches provide reasonable approximations.
In this chapter, I propose a more direct approach for estimating transition probabilitiesone designed for cross sectional analysis. This approach takes advantage of the fact that repeated cross sectional data sets, like the Current Population Survey (U.S.) and the Labour Force Survey (Canada), are representative of the country's population.

I propose two estimators of the transition probability, and clearly identify their identifying assumptions and asymptotic properties. I also show that existing cross sectional methods for estimating standard errors are downward biased and can lead to the identification of spurious breaks in job stability.

The structure of this section is as follows: Section 2.2 summarizes the job stability literature and Section 2.3 presents an overview of the empirical approach used in this thesis. Section 2.4 presents an alternative statistical representation of the retention rate - one better suited for cross sectional analysis. Section 2.5 offers an estimator for repeated cross sectional data. The asymptotic properties of the estimator are derived and a variance estimator proposed. Section 2.6 provides a discussion of the panel estimator. In Section 2.7, I contrast the panel and cross sectional estimators, and show how to apply the cross sectional estimator to survey data in Section 2.8. Section 2.9 provides a discussion of the bias of existing methods. In Section 2.10, I show how to test for changes in job stability, and, using CPS data, show that the choice of estimator can make a difference at the inference stage. I conclude with a discussion of how to construct an averaged estimator.

### 2.2 The Literature

The current literature on job duration and stability has its roots in the work of Hall (1982). Hall examined the prevalence of lifetime employment in the United States labour market, concluding that a large proportion of males had long employment spells, with no significant difference between whites and blacks. He also found the duration of employment for women shorter than for men.

Since the Current Population Survey (CPS) only measures in-progress job spells, and
not completed job duration, examining the distribution of in-progress jobs would underestimate the importance of long term employment. ${ }^{2}$ Hall was neither the first to recognize this difficulty nor to make an adjustment, but his introduction of a sound statistical framework to address those shortcomings was his main contribution. ${ }^{3}$ Specifically, he identified the importance of retention rates and their one-to-one mapping with the distribution of completed job spells. ${ }^{4}$

Formally, the $m$-period retention rate for workers with characteristics $c$ in year $t$ is simply the probability that such a worker remains with the same employer for an additional $m$ periods. Hall (1982) identified two methods of retention rate estimation: a historical approach which uses two years of data; and a contemporaneous one which only uses one cross section. An example will best illustrate the differences between the two approaches: to estimate the 10 -year retention rate for 40 year old workers with 5 years of job tenure in year $t$, the historical retention estimator takes the form
number of indiv. aged 50 with 15 years of tenure in the year $t+10$ sample
number of indiv. aged 40 with 5 years of tenure in the year $t$ sample

This estimator uses synthetic cohorts. It uses a group with similar characteristics, individuals that are now 50 years old in the period $t+10$ cross section, to estimate the number of workers from the denominator of equation (2.1) that would have stayed with the same employer for another 10 years. The construction of a complete job duration distribution would require a full set of historical retention rates, something not possible with infrequent data. ${ }^{5}$ Consequently, Hall advocated the contemporaneous method, requiring only a single

[^2]cross section, where you replace the numerator in equation (2.1) with an estimate from the year $t$ sample, and adjust for differences in population by age. ${ }^{6}$

Ureta (1992) re-examined Hall's work and came to different conclusions. She showed the contemporaneous approach requires constant job inflows over time, an assumption not supported in the data. ${ }^{7}$ Ureta found black men had a significantly lower probability of lifetime employment compared to white men. Women's job duration, however, was not as low as originally believed. Ureta faced the same identification problem as Hall; with data more than one year apart, one cannot directly calculate a full set of historical retention rates without making strong identifying assumptions. Instead of assuming constant job inflows, Ureta chose to assume a stable survival function. Diebold, Neumark, and Polsky (1997) later remarked that job stability should be tested, not assumed. Ureta was also criticized for her handling of the data. For supplements up to and including 1981, the question was "When did . . . start working at his present job or business?". Starting in 1983, the question became "How long has . . . been working continuously for his present employer (or self-employed)?". Although Ureta recognized the change, she did not (or could not) account for it in the analysis.

Starting in the late 1980s and early 1990s, much was made in the press of the disappearance of lifetime jobs (e.g. "Displaced Workers", Time, March 29, 1993). Reports of massive layoffs at major U.S. corporations were regular headline news. There was a common perception brought forth by the news media of workers becoming disposable, like any other resource or good. This led to renewed economic research on job duration, but with a slight

[^3]change in focus. For these second generation studies, job security and stability became the focal point of the research and not the job duration distribution itself. ${ }^{8}$

In an overview of the recent U.S. literature, Neumark (2000) found a consensus that no large scale decrease in job stability in the 1980s and 1990s ever happened. In the first wave of studies, there was some disagreement, with the choice of data set appearing to make a difference. Studies using the Panel Study of Income Dynamics (PSID) tended to find more significant changes in job stability compared to those using the CPS. The studies which used the CPS supplements as repeated cross-sections found little, if any, evidence of an overall decline in job stability during the same period.

Farber (1998) relied on the Displaced Worker Survey (DWS) supplements. ${ }^{9}$ He found that job stability tended to be cyclical, declining in periods of recessions, and increasing in periods of expansion. His results are not easily comparable to other studies since he did not use retention rates. Farber used a job loss rate, the ratio of the number of workers that have lost a job in the three years prior to the survey date to the number of workers at the time of the survey date. Unlike retention rates, there is no one-to-one mapping between the job loss rate and job stability. Additionally, the DWS supplement does not match up very well with other data sources. The DWS focussed on job displacement, and was not designed to examine other types of job separation, such as quits. The DWS would also count as a job loss a person that was laid off from a job and rehired in a different position with the same employer.

The remaining CPS based literature relies upon tenure supplements. There was initially some disagreement on the importance of the overall decline in job stability. This led to a debate between two groups of researchers: the group of Swinnerton and Wial who believed that there was an important decline, versus the group of Diebold, Hansen, Neumark and Polsky, who did not. The gap in estimates eventually narrowed for two reasons. First, Diebold, Neumark, and Polsky (1997) found that Swinnerton and Wial (1995) were not

[^4]consistent in their definition of the self-employed across supplements, a point later acknowledged by Swinnerton and Wial (1996). Once corrected, their results became more modest. Second, the business cycle adjustment used by Diebold, Neumark, and Polsky (1997) was put into question. Both job layoffs and quits had cyclical dimensions. Quits tended to be pro-cyclical, layoffs counter-cyclical. There was no clear a priori pattern to job separation and, therefore, no obvious adjustment to make. In the end, the estimated overall decline, when comparing the 1983-1987 and 1987-1991 periods, ranged from $2.2 \%$ (Neumark, Polsky, and Hansen 2000) to $3.5 \%$ (Swinnerton and Wial 1996). Over this same period, Neumark, Polsky, and Hansen (2000) found significant decreases in job stability for the less educated (high school or less), blacks and young workers (16-24 years of age). ${ }^{10}$ There was also a decrease in male job stability, a drop not present for women. Finally, workers with 2 to 9 years of tenure experienced a decline in stability, in contrast to low tenured (less than 2 years) and high tenured (more than 15 years) workers who experienced small increases. Because the comparison periods are at different stages of the business cycle, there is no easy interpretation for these results. The NBER identifies a trough in November 1982 and March 1991, implying a comparison of an expansionary period (1983-1987) with another period (1987-1991) whose tail end includes a recession. ${ }^{11}$

When the data were extended to the mid 1990s, the patterns changed. Neumark, Polsky, and Hansen (2000) found the decline in the overall rate was partially reversed; there were no further drops for males, blacks or young workers, and job stability for women had increased. The most dramatic change occurred at low tenure levels. Comparing the 19831987 and 1991-1995 periods, job stability for workers with initial tenure of less than two years increased by 5.3 percentage points. Yet, the 1990s data introduced new difficulties. The five year interval between 1991 and 1996 tenure supplements would generate a five year retention rate, not directly comparable to previous four year rates. Neumark, Polsky,

[^5]and Hansen (2000) were forced to rely on the February 1995 Contingent Work Supplement. This supplement focussed on workers in contingent jobs and other alternative employment arrangements. ${ }^{12}$ The Contingent Work supplement asked a different tenure question to those who were classified as contingent workers, and more importantly, the word "continuous" was eliminated. Using several tenure related questions from the 1996 supplement, Neumark, Polsky, and Hansen (2000) were able make adjustment to the 1995 data, improving its compatibility. It must be noted, nonetheless, that a change in supplement was necessary.

Farber (1998) found similar overall results for the 1991-1995 period, but he also faced some additional data problems. There was a change made to the core question in 1994, resulting in a significant group of individuals identifying themselves as displaced, but who were actually voluntarily separated from their job. Farber made a crude adjustment, and this dramatically altered his results. ${ }^{13}$ The data problems of the DWS, combined with the difficulties associated with a distributional approach, make it very difficult to draw conclusions from his research.

Before moving on to the examination of panel studies, I diverge from the examination of data difficulties associated with American cross sectional data, and focus on the fundamental goal of these second generation cross sectional studies-determining whether job stability has changed over time. The common approach in the literature has been to use a t-test to test for changes in job stability. Such an approach requires a clear understanding of both the asymptotic properties and the identifying assumptions of the estimator. In reality, neither is the case. To illustrate this, I must first discuss how the retention rate estimator, as presented in equation (2.1), is applied to survey data.

Cross sectional researchers take advantage of the fact that base weights of representative cross sections, like the Labour Force Survey (Canada) and the Current Population Survey (US), sum up to their respective populations. They replace the numerator in equation (2.1) with the sum of the base weights of all individuals 50 year old of age with 15 years of job

[^6]tenure in year $t+10$ (a similar procedure is used for the denominator). This weighted count (or sum) directly estimates the number of individual in the population aged 50 that have 15 years of tenure in period $t+10 .{ }^{14}$ By directly estimating a population count, this approach abstracts from difficulties associated with sample sizes changing over time-something not possible in equation (2.1). ${ }^{15}$

Although it may be intuitive that a larger sample will provide a better approximation of a population count, one cannot provide precision to that claim. Therefore, consistency can only be asserted, not proven. Finally, without a detailed proof of consistency, one cannot lay bare all underlying identifying assumptions. ${ }^{16}$

This lack of precision also carries over into the construction of standard errors. Two approaches have been followed in the literature. The first approach (e.g. Neumark, Polsky, and Hansen (1999); Swinnerton and Wial (1995)) applies an estimator designed for panel data to the cross section case, which fails to account for the variability resulting from an inability to follow individuals over time. The second approach proposed by Neumark, Polsky, and Hansen (2000) treats the numerator of the retention rate estimator as a random variable, but the denominator as a constant. ${ }^{17}$ Different draws from the year 1 population would generate different values for the denominator, and therefore, it should also be treated as random. I show in Section 2.10 that not accounting for the full variability can make a difference at the inference stage. More precisely, it may lead the researcher to falsely reject the null hypothesis of no change in job stability.

Neumark (2000) provides an interesting overview of PSID studies. The PSID, like the CPS, is a long lasting survey, beginning in 1968. Unlike the CPS, it has a panel

[^7]structure. ${ }^{18}$ The first wave of studies found significant declines in job stability over the 1970s and 1980s. Rose (1995) found the proportion of strong stayers, those with less than one job change, declined by $15 \%$ over this period. Marcotte (1995), examined data from 1976 through 1988, concluding that retention rates for whites and blacks fell by $8 \%$ and $13 \%$, respectively. These strong declines were in direct contrast to the CPS studies which found more modest results. Researchers (e.g., Diebold, Neumark, and Polsky (1997); Polsky (1999); and Jaeger and Stevens (2000)) that tried to reconcile the CPS and PSID results concluded that the problems associated with the tenure related questions of the PSID were at the root of the differences. The PSID does not have an employer ID code, which prevents the unique identification of an employment spell. In addition, the PSID has modified its tenure related questions every few years, including some significant changes over the 1984 to 1987 interval. ${ }^{19}$ Results varied dramatically depending on how the researcher dealt with these difficulties. The timing of the latest question revisions were particularly problematic. They occurred at the same time as the supposed changes in job stability.

Finally, the National Longitudinal Surveys (NLS) and the Survey of Income and Program Participation (SIPP), both panel data sets, have also been used in the job stability literature. Using the NLS, Bernhardt, Morris, Handcock, and Scott (1999) compared two panels of young white men and found some evidence of increased job instability in the 1980s and 1990s. By using two cohorts, the authors attempted to separate the cohort effect from the period effect. There are two main difficulties which hampers the generalization of their results. First, the experiences of the two cohorts are very different. For example, the younger cohort experienced the Vietnam War, while the other did not. More seriously, there are differences in the survey methods across cohorts (see Neumark (2000)). Gottschalk and Moffitt (1999) using the SIPP, find no evidence of increased job instability over the 1980s and 1990 s . As is the case for the NLS, the SIPP has a job identification code which can

[^8](in theory) lead to a 'cleaner' measure of job separation. The SIPP has two important advantages over other U.S. panels. It has overlapping panels which are interviewed every four months, in contrast with other U.S. panels which have annual or biannual interviews, and it interviews more households. Over the 1983-1993 period, the SIPP panels ranged from 12,527 to 23,627 households, and by 1996 had increased to over 40,000 households. The Bureau of Labour Statistics (BLS) has found some important inconsistencies with the job identification number across waves in the SIPP (see Stinson (2003)). Using administrative data sets, the BLS was able to provide a revised public use version for 1990-1993, but not for the 1980s.

Contrary to the United States, there have been few Canadian studies on job stability. Only Green and Riddell (1997) and Heisz $(1996,2005)$ have directly examined the long term changes in job stability across time in Canada. Both have used the Labour Force Survey (LFS), a rich source of tenure data. A systematic job tenure question has been part of the the monthly Labour Force Survey starting in 1976.

Green and Riddell (1997), using data from the public access March LFS files for 19791989, and 1991, found a hollowing of the middle of the job tenure distribution, with greater probability of short and long jobs. In the public access file, the tenure variable, including other important demographic characteristics, are only available by broad categories, making it impossible to directly calculate retention rates. As a result, Green and Riddell (1997) had to examine the distribution of in-progress jobs, a method more sensitive to job inflows.

Heisz (1996), who accessed to the restricted LFS files, also used a distribution approach to examine job stability and found similar results to Green and Riddell (1997). However, when Heisz (2005) extended the data through to 2001, the results became more complicated. Using both a distribution and a retention rate approach, he identified two phases for short term jobs: 1) job stability fell during the late 1970s until 1993, and then, 2) this trend reversed itself during the 1993-2001 period. He concluded that the changes experienced in the 1990s were only cyclical in nature. ${ }^{20}$ Finally, Heisz (2005) mentioned ageing of

[^9]the population as a possible explanation for some of his results, but did not explore this possibility in great detail. The focus of his study was to compare job stability between Canada and the United States, and as such was more descriptive in nature. An important goal of this thesis is to identify the source of the new patterns that are documented in this chapter.

### 2.3 Empirical Approach

This thesis uses a retention approach to identify the new job stability patterns in Canada: One advantage of this approach includes the direct link between job stability and retention rates - a job with the same employer is less stable if it has a lower probability of lasting one more period: Another important advantage is that a well-conditioned retention rate will be less sensitive to job inflows than in-progress job spell measures. Historically, Canada has experienced a large increase in labour force participation of women and a significant demographic change brought about by the baby-boomers, indicating that job inflows are an important issue for Canada.

The use of a retention rate approach is not novel; for reasons mentioned above, it has been the approach of choice in the job stability literature. What is unique to this thesis is how to apply the retention rate approach to cross sectional data. Unlike existing methods which adapt panel tools, I propose a method specifically designed for cross sectional data. The proposed method takes advantage of the fact that repeated cross sectional data sets, like the Current Population Survey (U.S.) and the Labour Force Survey (Canada), are representative of the country's population.

### 2.4 Population Retention Rate

The approach to econometric modeling in this thesis follows that of Goldberger (1991) and Wooldridge (2002), where the initial focus is the population and its moments. Assume a population distribution $F\left(x_{1}, x_{2}\right)$ where $x_{j}$ is a vector of characteristics for year $j$ such as age, gender, and tenure. This representation assumes a constant population, i.e for each
individual in the population there are two years of information. In this thesis, the year 1 population consists of all individuals living in Canada aged 20 to 54 . The year 2 population consists of the same individuals now one year older, i.e. aged 21 to $55 .{ }^{21}$

A one year retention rate for workers $a$ to $a^{\prime}$ years of age with $t$ to $t^{\prime}$ months of tenure is simply the probability that such a worker remains with the same employer for an additional year. ${ }^{22}$ Let $y_{i j}^{a, a^{\prime} ;, t, t^{\prime}}$ be a dummy variable which takes the value one if individual i is between $a$ and $a^{\prime}$ years of age and has between $t$ and $t^{\prime}$ months of tenure in period $j$. Use the shorthand notation $D_{i j}=y_{i j}^{a, a^{\prime} ; t, t^{\prime}}$ and $N_{i j}=y_{i j}^{a+1, a^{\prime}+1 ; t+12, t^{\prime}+12}$. The retention rate for a randomly chosen worker aged $a$ to $a^{\prime}$ with $t$ to $t^{\prime}$ months of tenure in year 1 is

$$
\begin{equation*}
R_{1}=R_{1}^{a, a^{\prime} ; t, t^{\prime}}=\operatorname{Prob}\left(N_{i 2}=1 \mid D_{i 1}=1\right) \tag{2.2}
\end{equation*}
$$

Given its conditional structure, equation (2.2) is a good starting point for panel data. One can condition on an individual working in the first period of a panel, and therefore estimate the sample analog of equation (2.2). With repeated cross sections, however, this is not possible. I propose an alternative representation of the population retention rate. Proposition 1 shows that it can be written as a function of two population moments, i.e.

$$
\begin{equation*}
R_{1}=\frac{E\left(N_{k 2}\right)}{E\left(D_{i 1}\right)} \tag{2.3}
\end{equation*}
$$

Proposition 1 Assuming a constant population over years 1 and 2, the population retention rate can be expressed as $R_{1}=\frac{E\left(N_{k 2}\right)}{E\left(D_{i 1}\right)}$.

[^10]proof:
\[

$$
\begin{aligned}
R_{1} & =\operatorname{Prob}\left(N_{i 2}=1 \mid D_{i 1}=1\right) \\
& =\frac{\operatorname{Prob}\left(N_{i 2}=1, D_{i 1}=1\right)}{\operatorname{Prob}\left(D_{i 1}=1\right)} \\
& \text { and since } N_{i 2}=1 \text { implies } D_{i 1}=1, \text { one can rewrite } R_{1} \text { as } \\
R_{1} & =\frac{\operatorname{Prob}\left(N_{i 2}=1\right)}{\operatorname{Prob}\left(D_{i 1}=1\right)} \\
& =\frac{E\left(N_{i 2}\right)}{E\left(D_{i 1}\right)} \\
& =\frac{E\left(N_{k 2}\right)}{E\left(D_{i 1}\right)}
\end{aligned}
$$
\]

Although cosmetic in nature, the introduction of the $k$ subscript in the last line of the proof is conceptually important. It emphasizes the fact the numerator of equation (2.3) is an unconditional moment; a moment that does not condition on period 1 events. A constant population is the key identifying assumption. In the next section, I argue that equation (2.3) is the appropriate initial step for cross sectional analysis.

### 2.5 Proposed Cross Sectional Estimator

With two cross sections, the proposed estimator is

$$
\begin{equation*}
\hat{R}_{1}=\hat{R}_{1}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\sum_{k=1}^{n_{2}} N_{k 2} / n_{2}}{\sum_{i=1}^{n_{1}} D_{i 1} / n_{1}} \tag{2.4}
\end{equation*}
$$

i.e. the sample analog of equation (2.3), where $n_{j}$ is the size of the year $j$ sample. ${ }^{23}$ Proposition 2 shows that the retention rate can be consistently estimated using two cross sections. Under clearly stated identifying assumptions, one can recover the retention rate without a panel data set.

Proposition 2 Assuming iid samples for each year, ${ }^{24}$ and no change in population, then

[^11]$\hat{R}_{1} \xrightarrow{p} R_{1}$
proof: Apply the Lindberg-Levy Central Limit Theorem
\[

$$
\begin{aligned}
& \sum_{k=1}^{n_{2}} N_{k 2} / n_{2} \xrightarrow{p} E\left(N_{k 2}\right) \\
& \sum_{i=1}^{n_{1}} D_{i 1} / n_{1} \xrightarrow{p} E\left(D_{i 1}\right)
\end{aligned}
$$
\]

and use the result of Proposition 1

$$
\frac{\sum_{k=1}^{n_{2}} N_{k 2} / n_{2}}{\sum_{i=1}^{n_{1}} D_{i 1} / n_{1}} \xrightarrow{p} \operatorname{Prob}\left(N_{i 2}=1 \mid D_{i 1}=1\right)
$$

Proposition 3 provides the asymptotic properties of the retention rate estimator in equation (2.4). I allow for both $N_{i j}$ and $D_{i j}$ to have sampling distributions. No restrictions are imposed on the correlation within observations, i.e. between $D_{i j}$ and $N_{i j}$, but independence across observations is assumed. For repeated cross sectional data this implies no correlation across time.

Proposition 3 Assuming iid samples for each year, no change in population, independence across years, and $\lim _{n_{1}, n_{2} \rightarrow \infty} \frac{n_{1}}{n_{2}}=1$, then $\sqrt{n}_{1}\left(\hat{R}_{1}-R_{1}\right) \xrightarrow{d} N(0, V)$ where $V$ is

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right) \tag{2.5}
\end{equation*}
$$

with

$$
\phi_{1}=\frac{E\left(N_{i 2}\right)}{\left[E\left(D_{i 1}\right)\right]^{2}}, \quad \phi_{2}=\frac{1}{E\left(D_{i 1}\right)}
$$

proof: For ease of notation let $\hat{N}_{j}=n_{j}^{-1} \sum_{i=1}^{n_{j}} N_{i j}, N_{j}=E\left(N_{i j}\right)$ and $V_{N_{j}}=V\left(N_{i j}\right)$, and define $\hat{D}_{j}, D_{j}$ and $V_{D_{j}}$ in a similar fashion.

$$
\begin{aligned}
\sqrt{n}_{1}\left(\hat{R}_{1}-R_{1}\right) & =\sqrt{n}_{1}\left(\frac{\hat{N}_{2}}{\hat{D}_{1}}-\frac{N_{2}}{D_{1}}\right) \\
& =\sqrt{n_{1}} \frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1} \hat{D}_{1}} \\
& =\frac{\sqrt{\frac{n_{1}}{n_{2}}} \sqrt{n_{2}}\left(\hat{N}_{2}-N_{2}\right) D_{1}-\sqrt{n}_{1}\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1}^{2}}+o_{p}(1) \\
& =-\phi_{1} \sqrt{n}_{1}\left(\hat{D}_{1}-D_{1}\right)+\phi_{2} \sqrt{n}_{2}\left(\hat{N}_{2}-N_{2}\right)+o_{p}(1) \\
& \xrightarrow{d} N\left(0, \phi_{1}^{2} V_{D_{1}}+\phi_{2}^{2} V_{N_{2}}\right)
\end{aligned}
$$

As equation (2.5) illustrates, $V$ is simply a weighted sum of the variance of $D_{i 1}$ and $N_{i 2}$, with the weights reflecting the non-linearity of the retention rate estimator. Replacing the population moments in (2.5) with corresponding sample analogs generates a consistent estimator of the asymptotic variance.

### 2.6 Panel Estimator

Assume a sample of characteristics of n individuals over two time periods. Without loss of generality, further assume the first $n^{a, a^{\prime} ; t, t^{\prime}}$ individuals are between $a$ to $a^{\prime}$ years of age and have between $t$ to $t^{\prime}$ months of tenure in year 1. For each individual in the sub-sample of size $n^{a, a^{\prime} ; t, t^{\prime}}$, let $X_{i}^{a, a^{\prime} ; t, t^{\prime}}$ be a dummy variable equal to 1 if individual $i$ is with the same employer in year 2 and zero otherwise. The mean of $X_{i}^{a, a^{\prime} ; t, t^{\prime}}$, i.e. $E\left(X_{i}^{a, a^{\prime} ; t, t^{\prime}}\right)$, is the population retention rate. As a result, the panel estimator proposed in the literature is

$$
\begin{equation*}
\hat{R P_{1}}=\hat{R P_{1}^{a, a^{\prime} ; t, t t^{\prime}}}=\frac{\sum_{i=1}^{n^{a, a^{\prime} ; t, t^{\prime}}} X_{i}^{a, a^{\prime} ; t, t t^{\prime}}}{n^{a, a^{\prime} ; t, t t^{\prime}}} \tag{2.6}
\end{equation*}
$$

i.e. the sample analog of equation (2.2). The known properties of this estimator are formally stated in Proposition 4.

Proposition 4 Assume an iid sample of characteristics of $n$ individuals over two time
period where the first $n^{a, a^{\prime} ; t, t^{\prime}}$ individuals are between a to $a^{\prime}$ years of age and have between. $t$ to $t^{\prime}$ months of tenure in year 1. Then, $\sqrt{n^{a, a ; t, t t^{\prime}}}\left(\hat{R P_{1}}-R_{1}\right) \xrightarrow{d} N(0, V)$ where $V$ is

$$
\begin{equation*}
V=R_{1}\left(1-R_{1}\right) \tag{2.7}
\end{equation*}
$$

proof: Apply the Lindberg-Levy Central Limit Theorem and use the fact that the variance of $X_{i}^{a, a^{\prime} ; t, t^{\prime}}$, a Bernoulli random variable, is $E\left(X_{i}^{a, a^{\prime} ; t, t^{\prime}}\right)\left(1-E\left(X_{i}^{a, a^{\prime} ; t, t^{\prime}}\right)\right)$.

I suggest an alternative representation for the panel estimator, the sample analog of

$$
\begin{equation*}
\hat{R P}_{1}=\hat{\hat{R P}}{ }_{1}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\sum_{i=1}^{n} N_{i 2} / n}{\sum_{i=1}^{n} D_{i 1} / n} \tag{2.3}
\end{equation*}
$$

$\hat{R P_{1}}$ generates the same estimates as $\hat{R P_{1}}$, and it can easily be shown that they share the same asymptotic properties. But by using the $N_{i j}$ and $D_{i j}$ notation, $\hat{R P_{1}}$ is more easily comparable with the cross sectional estimator, $\hat{R}_{1}$. Both panel estimator, $\hat{R P_{1}}$, and cross sectional estimator, $\hat{R}_{1}$, are ratios of sample means, and as such, have similar variance structure. There is an important difference. $D_{i 1}$ and $N_{i 2}$ are correlated in a panel, not in repeated cross sections. As a result, the asymptotic variance of the panel estimator, $\hat{R P_{1}}$, will require an additional covariance term. The properties of this estimator are provided in Proposition 5.

Proposition 5 Assume an iid sample of characteristics of $n$ individuals over two time period, then $\sqrt{n}\left(\hat{\hat{R P}_{1}}-R_{1}\right) \xrightarrow{d} N(0, V)$ where $V$ is

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right)-2 \phi_{1} \phi_{2} \operatorname{Cov}\left(D_{i 1}, N_{i 2}\right) \tag{2.9}
\end{equation*}
$$

with

$$
\phi_{1}=\frac{E\left(N_{i 2}\right)}{\left[E\left(D_{i 1}\right)\right]^{2}}, \quad \phi_{2}=\frac{1}{E\left(D_{i 1}\right)}
$$

proof: For ease of notation let $\hat{N}_{j}=n^{-1} \sum_{i=1}^{n} N_{i j}, N_{j}=E\left(N_{i j}\right)$ and $V_{N_{j}}=V\left(N_{i j}\right)$, and define $\hat{D}_{j}, D_{j}$ and $V_{D_{j}}$ in a similar fashion. Finally, let $C_{2}=\operatorname{Cov}\left(D_{i 1}, N_{i 2}\right)$.

$$
\begin{aligned}
\sqrt{n}\left(\hat{R P P}_{1}-R_{1}\right) & =\sqrt{n}\left(\frac{\hat{N}_{2}}{\hat{D}_{1}}-\frac{N_{2}}{D_{1}}\right) \\
& =\sqrt{n} \frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1} \hat{D}_{1}} \\
& =\frac{\sqrt{n}\left(\hat{N}_{2}-N_{2}\right) D_{1}-\sqrt{n}\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1}^{2}}+o_{p}(1) \\
& =-\phi_{1} \sqrt{n}\left(\hat{D}_{1}-D_{1}\right)+\phi_{2} \sqrt{n}\left(\hat{N}_{2}-N_{2}\right)+o_{p}(1) \\
& \xrightarrow{d} N\left(0, \phi_{1}^{2} V_{D_{1}}+\phi_{2}^{2} V_{N_{2}}-2 \phi_{1} \phi_{2} C_{2}\right)
\end{aligned}
$$

### 2.7 Comparison of Panel and Cross Sectional Estimator

I have shown that both panel and cross sectional estimators can consistently estimate the population retention rate. Given the asymptotic properties presented in Section 2.5 and 2.6 , one can also construct t-tests (to test for changes in job stability across time and group) using either approach. ${ }^{25}$ I argue that the cost of not following an individual over time is only one of efficiency. Assuming a two-year panel, and no attrition, satisfies the identifying assumptions of the panel estimator (see Proposition 5) and cross sectional estimator (see Proposition 3). As a result, the asymptotic variance of the panel estimator can be written as

$$
\begin{equation*}
\operatorname{Avar}\left(\hat{R P}_{1}\right)=\operatorname{Avar}\left(\hat{R}_{1}\right)+\left(-\frac{E\left(N_{i 2}\right)}{\left[E\left(D_{i 1}\right)\right]^{3}} \operatorname{Cov}\left(D_{i 1}, N_{i 2}\right)\right) \tag{2.10}
\end{equation*}
$$

i.e. the asymptotic variance of the cross sectional estimator plus a negative term. The second righthand side term is negative since $D_{i 1}$ and $N_{i 2}$ are positively correlated. This negative term reflects the efficiency gains of being able to follow an individual over time in a panel data set.

[^12]
### 2.8 Application to Survey Data

The LFS has non-random aspects to its design, and as a result, the probability of being selected in each survey is not the same across observations. The standard solution is to use weights. The new retention rate estimator is

$$
\begin{equation*}
\hat{\hat{R}}_{1}=\hat{\hat{R}}_{1}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\sum_{k=1}^{n_{2}} n w_{k 2} N_{k 2} / n_{2}}{\sum_{i=1}^{n_{1}} n w_{i 1} D_{i 1} / n_{1}} \tag{2.11}
\end{equation*}
$$

where $n w_{i j}$ is the normalized weight for individual $i$ in year $j$ (i.e. $n w_{i j}=\frac{b w_{i j}}{\sum_{i=1}^{n_{j} b w_{i j} / n_{j}}}$ and where $b w$ is the base weight provided by the LFS). Relaxing the non-randomness assumption does not alter the asymptotic results, i.e. $\hat{\hat{R}}_{1}$ will have the same asymptotic properties as $\hat{R}_{1}$. For the variance estimator, one replaces the population moments in (2.5) with the weighted sample analogs (using the normalized weights).

More importantly, the population has not remained constant over time. Changes in population due to deaths, emigration and immigration, will bias the results. ${ }^{26}$

Given that the sum of the sample's base weights is Canada's target population, ${ }^{27}$ I propose an alternative estimator which can address population changes

$$
\begin{equation*}
\tilde{R}_{1}=a d j_{1} \hat{\hat{R}}_{1} \tag{2.12}
\end{equation*}
$$

where $a d j_{1}=\frac{\sum_{i=1}^{n_{2}} b w_{i 2}}{\sum_{i=1}^{n 1} b w_{i 1}}$. If the population did not change, the adjustment factor, i.e. $a d j_{1}$, would be equal to 1 , and the two estimators, i.e. $\hat{\hat{R}}_{1}$ and $\tilde{R}_{1}$, would be numerically equivalent. ${ }^{28}$ With a population change, the estimator $\tilde{R}_{1}$ will consistently estimate the true retention rate if the following conditions hold: the population changes are due to immigration and emigration, and that these migration decisions lead to breaks in tenure spell. ${ }^{29}$

[^13]This estimator shares the same asymptotic properties as the others, but the variance estimator needs adjustment to account for the population change, i.e. not following a population cohort. Treating $a d j_{1}$ as a constant, the variance estimator is the (weighted) sample analogs of the population moments in equation (2.5), pre-multiplied by $\left(a d j_{1}\right)^{2}$.

The empirical results of this paper are based on $\tilde{R}_{1}$ for the following reasons. Immigration has played an important role in the population growth in Canada, and the potential importance of death as a source of population change is reduced by limiting the target population in year 2 to individuals aged 21 to 55 . Also, stating that migration affects tenure does allow for immigrant employment in year 2. It only assumes new immigrants do not keep the same employer when moving to Canada. ${ }^{30}$ Finally, there is an issue of comparability. This paper attempts to shed light on the changes in job stability not only in Canada, but also in the United States. Due to data limitation, $\tilde{R}_{1}$ is the preferred estimator for U.S. retention rates. Tenure in the CPS is only collected in select supplements, and as a result, the U.S. analysis is restricted to 4 year retention rates. Using $\hat{\hat{R}}_{1}$ would implicitly assume that the population had not changed over the 4 year interval, a much too restrictive assumption.

Table 2.1: LFS Adjustment factor $\left(a d j_{j}\right)$ by Year

| Year | Adjustment <br> factor | Year | Adjustment <br> factor | Year <br> Year | Adjustment <br> factor |
| :--- | :---: | :---: | :---: | :---: | :---: |
| 1977 | 1.0008 | 1986 | 0.9973 | 1995 | 1.0062 |
| 1978 | 1.0013 | 1987 | 1.0057 | 1996 | 1.0017 |
| 1979 | 1.0033 | 1988 | 1.0076 | 1997 | 0.9972 |
| 1980 | 1.0022 | 1989 | 1.0017 | 1998 | 1.0028 |
| 1981 | 1.0021 | 1990 | 1.0043 | 1999 | 1.0067 |
| 1982 | 0.9983 | 1991 | 1.0052 | 2000 | 1.0076 |
| 1983 | 1.0014 | 1992 | 1.0047 | 2001 | 1.0071 |
| 1984 | 0.9997 | 1993 | 1.0063 | 2002 | 1.0052 |
| 1985 | 1.0014 | 1994 | 1.0030 | 2003 | 1.0092 |

Table 2.1 shows the adjustment factor, i.e. $a d j_{j}$, for 1977 through 2003. That factor is very close to 1 for all years, and as a result the $\tilde{R}_{1}$ and $\hat{\hat{R}}_{1}$ estimates (and their respective
rate should be

$$
R_{1}=\left(\frac{n+1}{n}\right) \frac{\sum_{i=1}^{n+1} N_{i 2} / n+1}{\sum_{i=1}^{n} D_{i 1} / n}
$$

${ }^{30} \mathrm{~A}$ similar argument follows for an emigrant.
variance estimates) will be very similar. This illustrates the strength of the LFS. Even with very different identifying assumptions, the frequency of the LFS tenure data ensures similar results.

### 2.9 Existing Cross Sectional Methods and their Biases

In this section I re-examine the cross sectional methods proposed in the literature. I compare and contrast the existing approach with those proposed in this thesis, effectively identifying the weaknesses and biases of existing methods.

The cross sectional literature uses the panel estimator of equation (2.6) as its starting point. Given two random samples of workers where $n_{1}^{*}$ and $n_{2}^{*}$ are the number of individuals that work in samples 1 and 2 , respectively, the cross sectional approximation is

$$
\begin{equation*}
\hat{Q}_{1}=\hat{Q}_{1}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\sum_{k=1}^{n^{*}} N_{k 2}}{\sum_{i=1}^{n^{*}} D_{i 1}} \tag{2.13}
\end{equation*}
$$

Recognizing that the estimator in equation (2.13) is sensitive to sample sizes, researchers have used the following weight-based method when faced with survey data.

$$
\begin{equation*}
\tilde{Q}_{1}=\tilde{Q}_{1}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\sum_{i=1}^{n_{2}^{*}} b w_{i 2} N_{i 2}}{\sum_{i=1}^{n_{1}^{*}} b w_{i 1} D_{i 1}} \tag{2.14}
\end{equation*}
$$

As previously discussed in Section (2.2), this estimator takes advantage of the fact that the sum of the base weights of repeated cross sectional data sets, like the LFS and CPS, add up to their target population. The denominator and numerator of equation (2.14) have a very natural interpretation. The denominator is an estimate of the number of Canadians aged $a$ to $a^{\prime}$ with $t$ to $t^{\prime}$ months of initial tenure in year 1 , and the denominator an estimate of the number of Canadians aged $a+1$ to $a^{\prime}+1$ that have $t+12$ and $t^{\prime}+12$ months of tenure in year 2 .

There are three important reasons why the proposed methods are preferable to the existing one - regardless of the fact that $\tilde{Q}_{1}$ and $\tilde{R}_{1}$ are numerically equivalent. ${ }^{31}$ These

[^14]three reasons are:
One, only the proposed methods offer clear identification of the assumptions necessary for consistency. The $\hat{\hat{R}}_{1}$ estimator assumes that the samples are drawn from a constant population. The $\tilde{R}_{1}$ estimator allows for population changes over time, but in a controlled way; it requires that population changes due to migration decisions lead to breaks in tenure spells. With the existing method, i.e. $\tilde{Q}_{1}$, the literature can only claim, but not prove, that a larger sample will provide a better approximation of the retention rate. Without such a proof, one cannot lay bare all underlying identifying assumptions. A clear understanding of the consistency requirements is critical considering that an important selling point of cross sectional data sets like the CPS and LFS have been their large sample sizes.

Two, the existing method treats base weights as counts-which is not a typical approach in empirical work. The standard empirical approach is to normalize the weights so that they sum up to the sample size. These normalized weights are then used to account for the fact that some observations had a higher probability to be drawn than others-as is done in $\hat{\hat{R}_{1}}$. In other words, the weights are used as relative measures. By using the weights as counts, $\tilde{Q}_{1}$ (and $\tilde{R}_{1}$ for that matter) imposes an additional identifying assumption: that the counts are accurate. This assumption some researchers may not be willing to make.

Three, the standard error estimators based on $\tilde{Q}_{1}$ are downward biased. Two approaches to the variance estimator have been used in the literature. ${ }^{32}$ The first approach (e.g., Diebold, Neumark, and Polsky (1997); Swinnerton and Wial (1995)) applies the variance estimator designed for longitudinal data to the repeated cross section case, i.e. they estimate $V$ in equation (2.7) by replacing $R_{1}$ with $\hat{Q}_{1}$. This estimator does not account for the variability resulting from the inability to follow individuals over time. Thus, the estimator ought to have a downward bias. Equation (2.10) confirms this intuition. As discussed in Section 2.7) the second (negative) righthand side term in equation (2.10) reflects the efficiency gains that result from following an individual over time in a panel-gains that cannot be had with repeated cross sections. By estimating equation (2.10), one is therefore

[^15]underestimating the true variance of the cross sectional estimator. The second approach, proposed by Neumark, Polsky, and Hansen (1999), allows $N_{k 2}$ in equation (2.13) to have a sampling distribution, but assumes that $D_{i 1}$ is a constant. The main difficulty with this approach is that different draws from the year 1 population would generate different values for $D_{i 1}$, and as such, it should also be treated as a random variable. Therefore, the Neumark, Polsky and Hansen (NPH) estimator only accounts for part of the variability introduced by the synthetic cohort approach. More precisely, the NPH variance estimator can only account for the first of two righthand side terms in equation (2.5). Conditioning on a sample distribution for $D_{i 1}$, a larger $\hat{N}_{2}$ is associated with a larger second righthand side term. As a result, the extent to which the NPH method underestimates the variance may be correlated with the size of the retention rate. ${ }^{33}$

In the next section, I show how to test for changes in job stability. In addition, I show that existing methods cannot properly account for covariance terms that arise when testing for changes across time or groups, introducing further bias at the inference stage.

### 2.10 Testing for Changes in Job Stability

Within the retention approach, testing for change in job stability across time or groups is straightforward-only a single restriction needs to be tested. For differences across time, one tests the null hypothesis $H_{0}: R_{j}-R_{1}=0$ against the alternative $H_{a}: R_{j}-R_{1} \neq 0$, where $R_{j}-R_{1}$ is the difference in retention rate over a $j-1$ period. For differences across groups, say $A$ and $B$, the null and alternative hypothesis are $H_{0}: R_{j}^{A}-R_{j}^{B}=0$ and $H_{a}: R_{j}^{A}-R_{j}^{B} \neq 0$, respectively.

I use a t-test to test these restrictions. ${ }^{34}$ The t-statistic for differences across time, $t_{n}$, takes the form

$$
\begin{equation*}
t_{n}=\frac{\hat{R}_{j}-\hat{R}_{1}}{\sqrt{\hat{V}_{R_{j}-R_{1}} / n}} \tag{2.15}
\end{equation*}
$$

[^16]where $\hat{V}_{R_{j}-R_{1}}$ is the estimator of $\operatorname{Avar}\left(\hat{R}_{j}-\hat{R}_{1}\right)$. For differences across groups, the test statistic, $g_{n}$, is
\[

$$
\begin{equation*}
g_{n}=\frac{\hat{R}_{j}^{A}-\hat{R}_{j}^{B}}{\sqrt{\hat{V}_{R_{j}^{A}-R_{j}^{B}} / n}} \tag{2.16}
\end{equation*}
$$

\]

where $\hat{V}_{R_{j}^{A}-R_{j}^{B}}$ is the estimator of $\operatorname{Avar}\left(\hat{R}_{j}^{A}-\hat{R}_{j}^{B}\right)$.
In both cases, the variance estimator must account for potential covariance terms. When testing whether job stability changed from one period to the next, one must account for the fact that $\hat{R}_{2}$ and $\hat{R}_{1}$ may be correlated since both the denominator of $\hat{R}_{2}$ and the numerator of $\hat{R}_{1}$ are functions of the same (year 2) observations. Similarly, the numerators (and denominators) of $R_{j}^{A}$ and $R_{j}^{B}$ may also be correlated.

Existing approaches to estimating standard errors cannot easily account account for these possibilities. The NPH method, for example, rules out the possibility of any correlation between $\hat{R}_{2}$ and $\hat{R}_{1}$ by assuming that $D_{i j}$ is a constant. By allowing both $N_{i j}$ and $D_{i j}$ to have sampling distributions, the method proposed in this thesis can account for potential covariance terms. Proposition 6 provides the asymptotic properties for the difference in retention rate over time. ${ }^{35}$

Proposition 6 Assuming iid samples for each year, samples of equal size, independence across years, and no change in population, then $\sqrt{n}\left(\left(\hat{R}_{j}-\hat{R}_{1}\right)-\left(R_{j}-R_{1}\right) \xrightarrow{d} N(0, V)\right.$ where $V$ depends on $j$, an integer greater than or equal to 2.

Case 1: $j=2$

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right)+\phi_{3}^{2} V\left(D_{i 2}\right)+\phi_{4}^{2} V\left(N_{i 3}\right)+2 \phi_{2} \phi_{3} \mu \operatorname{Cov}\left(D_{i 2}, N_{i 2}\right) \tag{2.17}
\end{equation*}
$$

Case 2: $j \geq 3$

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right)+\phi_{3}^{2} V\left(D_{i j}\right)+\phi_{4}^{2} V\left(N_{i j+1}\right) \tag{2.18}
\end{equation*}
$$

[^17]with
$$
\phi_{1}=\frac{E\left(N_{i 2}\right)}{\left[E\left(D_{i 1}\right)\right]^{2}}, \quad \phi_{2}=\frac{1}{E\left(D_{i 1}\right)}, \quad \dot{\phi}_{3}=\frac{E\left(N_{i j+1}\right)}{\left[E\left(D_{i j}\right)\right]^{2}}, \quad \phi_{4}=\frac{1}{E\left(D_{i j}\right)}
$$
and $\mu$ is the probability that a random chosen person in the population aged 20 to 54 will be 21 to 54 .
proof: For ease of notation let $\hat{N}_{j}=n_{j}^{-1} \sum_{i=1}^{n_{j}} N_{i j}, N_{j}=E\left(N_{i j}\right)$ and $V_{N_{j}}=V\left(N_{i j}\right)$, and define $\hat{D}_{j}, D_{j}$ and $V_{D_{j}}$ in a similar fashion. Finally, let $C_{2}=\operatorname{Cov}\left(D_{i 2}, N_{i 2}\right)$

Case 1: $j=2$

$$
\begin{aligned}
& \sqrt{n}\left(\left(\hat{R}_{2}-\hat{R}_{1}\right)-\left(R_{2}-R_{1}\right)\right. \\
& =\sqrt{n}\left(\left(\frac{\hat{N}_{3}}{\hat{D}_{2}}-\frac{N_{3}}{D_{2}}\right)-\left(\frac{\hat{N}_{2}}{\hat{D}_{1}}-\frac{N_{2}}{D_{1}}\right)\right) \\
& =\sqrt{n} \frac{\left(\hat{N}_{3}-N_{3}\right) D_{2}-\left(\hat{D}_{2}-D_{2}\right) N_{3}}{D_{2} \hat{D}_{2}}-\sqrt{n} \frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1} \hat{D}_{1}} \\
& =\sqrt{n} \frac{\left(\hat{N}_{3}-N_{3}\right) D_{2}-\left(\hat{D}_{2}-D_{2}\right) N_{3}}{D_{2}^{2}}-\sqrt{n} \frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1}^{2}}+o_{p}(1) \\
& =\phi_{1} \sqrt{n}\left(\hat{D}_{1}-D_{1}\right)-\phi_{2} \sqrt{n}\left(\hat{N}_{2}-N_{2}\right)-\phi_{3} \sqrt{n}\left(\hat{D}_{2}-D_{2}\right)+\phi_{4} \sqrt{n}\left(\hat{N}_{3}-N_{3}\right)+o_{p}(1) \\
& \xrightarrow{d} N\left(0, \phi_{1}^{2} V_{D_{1}}+\phi_{2}^{2} V_{N_{2}}+\phi_{3}^{2} V_{D_{2}}+\phi_{4}^{2} V_{N_{3}}+2 \phi_{2} \phi_{3} \mu C_{2}\right)
\end{aligned}
$$

Case 2: $j \geq 3$. The proof is similar to Case 1 , with one exception. Since the four components of the test statistics, i.e. $\hat{N}_{j+1}, \hat{D}_{j}, \hat{N}_{2}$ and $\hat{D}_{1}$ are functions of different yearly samples when $j \geq 3$, the covariance term is zero.

Replacing the population moments in (2.17) and (2.18) with corresponding sample analogs generates a consistent estimator for each asymptotic variance.

The flexibility of the proposed approach is apparent when testing for a change in job stability over consecutive periods. The fifth term in (2.17) represents the covariance between $\hat{R}_{1}$ and $\hat{R}_{2}$; a term whose sign depends on the sign of $\operatorname{Cov}\left(N_{i 2}, D_{i 2}\right)$. Neumark, Polsky, and Hansen (1999) claim the covariance between $\hat{R}_{1}$ and $\hat{R}_{2}$ will be positive if the probability that both $N_{i 2}$ and $D_{i 2}$ equal one is also positive. In fact, the covariance term will be positive
if and only if ${ }^{36}$

$$
\begin{equation*}
\operatorname{Pr}\left(N_{i 2}=1, D_{i 2}=1\right)>\operatorname{Pr}\left(N_{i 2}=1\right) \operatorname{Pr}\left(D_{i 2}=1\right) \tag{2.19}
\end{equation*}
$$

The cases of $t^{\prime}=t+1 / 2$ versus $t^{\prime}=\infty$ illustrate how a change in the retention rate's conditioning characteristics can alter the sign of the covariance. If the retention rate conditions on a narrowly defined tenure interval, i.e. $t^{\prime}=t+1 / 2$, the probability that both $D_{i j}$ and $N_{i j}$ equal one is zero, and thus, $\operatorname{Cov}\left(N_{i 2}, D_{i 2}\right)<0$. If instead, the retention rate conditions on a very large tenure interval, i.e. $t^{\prime}=\infty$, then $P\left(N_{i 2}=1, D_{i 2}=1\right)=P\left(N_{i 2}=1\right)$, and as such, $\operatorname{Cov}\left(N_{i 2}, D_{i 2}\right)>0$. For most cases, though, it is not clear, a priori, whether equation (2.19) holds. As a result, ignoring the covariance term can under- or overstate the standard errors used in the t-statistic.

Proposition 7 provides the asymptotic properties for the difference in retention rate across groups.

Proposition 7 Assuming iid samples for each year, samples of equal size, independence across years, and no change in population, then $\sqrt{n}\left(\left(\hat{R}_{1}^{A}-\hat{R}_{1}^{B}\right)-\left(R_{1}^{A}-R_{1}^{B}\right) \xrightarrow{d} N(0, V)\right.$ where $V$ depends on $j$, an integer greater than or equal to 2.

$$
\begin{align*}
V= & \phi_{1}^{2} V\left(D_{i 1}^{A}\right)+\phi_{2}^{2} V\left(N_{i 2}^{A}\right)+\phi_{3}^{2} V\left(D_{i 1}^{B}\right)+\phi_{4}^{2} V\left(N_{i 2}^{B}\right) \\
& -2 \phi_{1} \phi_{3} \operatorname{Cov}\left(D_{i 1}^{A}, D_{i 1}^{B}\right)-2 \phi_{2} \phi_{4} \operatorname{Cov}\left(N_{i 2}^{A}, N_{i 2}^{B}\right) \tag{2.20}
\end{align*}
$$

with

$$
\phi_{1}=\frac{E\left(N_{i 2}^{A}\right)}{\left[E\left(D_{i 1}^{A}\right)\right]^{2}}, \quad \phi_{2}=\frac{1}{E\left(D_{i 1}^{A}\right)}, \quad \phi_{3}=\frac{E\left(N_{i 2}^{B}\right)}{\left[E\left(D_{i 1}^{B}\right)\right]^{2}}, \quad \phi_{4}=\frac{1}{E\left(D_{i 1}^{B}\right)}
$$

proof: For ease of notation let $\hat{N}_{j}^{K}=n^{-1} \sum_{i=1}^{n} N_{i j}^{K}, N_{j}^{K}=E\left(N_{i j}^{K}\right)$ and $V_{N_{j}^{K}}=V\left(N_{i j}^{K}\right)$ for $\mathrm{K}=\mathrm{A}, \mathrm{B}$, and define $\hat{D}_{j}^{K}, D_{j}^{K}$ and $V_{D_{j}^{K}}$ in a similar fashion. Finally, let $C_{1}=\operatorname{Cov}\left(D_{i 1}^{A}, D_{i 1}^{B}\right)$

[^18]and $C_{2}=\operatorname{Cov}\left(D_{i 2}^{A}, D_{i 2}^{B}\right)$
\[

$$
\begin{aligned}
& \sqrt{n}\left(\left(\hat{R}_{1}^{A}-\hat{R}_{1}^{B}\right)-\left(R_{1}^{A}-R_{1}^{B}\right)\right. \\
& =\sqrt{n}\left(\left(\frac{\hat{N}_{2}^{A}}{\hat{D}_{1}^{A}}-\frac{N_{2}^{B}}{D_{1}^{B}}\right)-\left(\frac{\hat{N}_{2}^{A}}{\hat{D}_{1}^{A}}-\frac{N_{2}^{B}}{D_{1}^{B}}\right)\right) \\
& =\sqrt{n} \frac{\left(\hat{N}_{2}^{A}-N_{2}^{A}\right) D_{1}^{A}-\left(\hat{D}_{1}^{A}-D_{1}^{A}\right) N_{2}^{A}}{D_{1}^{A} \hat{D}_{1}^{A}}-\sqrt{n} \frac{\left(\hat{N}_{2}^{B}-N_{2}^{B}\right) D_{1}^{B}-\left(\hat{D}_{1}^{B}-D_{1}^{B}\right) N_{2}^{B}}{D_{1}^{B} \hat{D}_{1}^{B}} \\
& =\sqrt{n} \frac{\left(\hat{N}_{2}^{A}-N_{2}^{A}\right) D_{1}^{A}-\left(\hat{D}_{1}^{A}-D_{1}^{A}\right) N_{2}^{A}}{\left(D_{2}^{A}\right)^{2}}-\sqrt{n} \frac{\left(\hat{N}_{2}^{B}-N_{2}^{B}\right) D_{1}^{B}-\left(\hat{D}_{1}^{B}-D_{1}^{B}\right) N_{2}^{B}}{\left(D_{1}^{B}\right)^{2}}+o_{p}(1) \\
& =-\phi_{1} \sqrt{n}\left(\hat{D}_{1}^{A}-D_{1}^{A}\right)+\phi_{2} \sqrt{n}\left(\hat{N}_{2}^{A}-N_{2}^{A}\right)-\phi_{3} \sqrt{n}\left(\hat{D}_{1}^{B}-D_{1}^{B}\right)+\phi_{4} \sqrt{n}\left(\hat{N}_{2}^{B}-N_{2}^{F}\right)+o_{p}(1) \\
& \xrightarrow{d} N\left(0, \phi_{1}^{2} V_{D_{1}^{A}}+\phi_{2}^{2} V_{N_{2}^{A}}+\phi_{3}^{2} V_{D_{1}^{B}}+\phi_{4}^{2} V_{N_{2}^{B}}-2 \phi_{1} \phi_{3} C_{1}-2 \phi_{2} \phi_{4} C_{2}\right)
\end{aligned}
$$
\]

Replacing the population moments in (2.20) with corresponding sample analogs generates a consistent estimator for the asymptotic variance. The fifth terms of equation (2.20) account for the correlation between the numerators of $\hat{R}_{1}^{A}$ and $\hat{R}_{1}^{B}$; a term whose signs depends on the sign of $-\operatorname{Cov}\left(N_{i 2}^{A}, N_{i 2}^{B}\right)$. Similarly, the sign of the sixth term will depend on the sign of $-\operatorname{Cov}\left(D_{i 1}^{A}, D_{i 1}^{B}\right)$. Contrasting two mutually exclusive groups, such as males and females, ensures that both covariance terms are negative. As a result, ignoring the covariance terms will understate the standard errors.

### 2.10.1 Application to CPS data

Using Current Population Survey data, I show that using the standard errors suggested in the literature (with their downward bias) may lead the researcher to falsely reject the null hypothesis of no change in job stability.

Tables 2.2 and 2.3 examine changes in 4-year retention rates from 1996 to 2000 for males and females, respectively. Standard errors are calculated for the NPH method, the DPP method, the proposed method, and the proposed method with no covariance term. In all cases, weights were used to make a clearer comparison of the various methods. ${ }^{37}$

[^19]Table 2.2: 4-year Male Retention Rates - Time Differentials

| Tenure Group Specification | 1996 | 2000 | Difference | Standard Errors Method |
| :---: | :---: | :---: | :---: | :---: |
| 0-2 | 0.4775 | 0.4999 | $\begin{gathered} 0.0224 \\ (0.0076)^{* *} \\ (0.0106)^{* *} \\ (0.0112)^{* *} \\ (0.0116)^{*} \end{gathered}$ | DNP <br> NPH <br> proposed proposed (no covariance) |
| 3-6 | 0.4522 | 0.4658 | $\begin{gathered} 0.0136 \\ (0.0092) \\ (0.0133) \\ (0.0147) \\ (0 ; 0150) \end{gathered}$ | DNP <br> NPH <br> proposed <br> proposed (no covariance term) |
| 7-11 | 0.7069 | 0.6554 | $\begin{gathered} -0.0515 \\ (0.0111)^{* *} \\ (0.0197)^{* *} \\ (0.0251)^{* *} \\ (0.0245)^{* *} \end{gathered}$ | DNP <br> NPH <br> proposed <br> proposed (no covariance term) |
| $12+$ | 0.7288 | 0.6927 | $\begin{gathered} -0.0362 \\ (0.0083)^{* *} \\ (0.0149)^{* *} \\ (0.0228) \\ (0.0193)^{*} \end{gathered}$ | DNP <br> NPH <br> proposed <br> proposed (no covariance term) |
| total | 0.5798 | 0.5717 | $\begin{gathered} -0.0081 \\ (0.0045)^{*} \\ (0.0060) \\ (0.0076) \\ (0.0067) \\ \hline \end{gathered}$ | DNP <br> NPH <br> proposed <br> proposed (no covariance term) |

[^20]Confirming the prediction of the theory, Tables 2.3 and 2.2 indicate that accounting for the covariance term either increases or decreases the standard errors, depending on the conditioning characteristics. Yet, even when the covariance term is negative, a consistent pattern emerges with the proposed method generating standard errors consistently larger than either DNP or NPH methods. ${ }^{38}$ My method generates standard errors that are up to $175 \%$ larger than the DNP estimates and up to $57 \%$ larger than the NPH estimates. The gap in standard errors is larger for higher age and tenure groups - both groups with higher

[^21]Table 2.3: 4-year Female Retention Rates - Time Differentials

| Tenure Group Specification | 1996 | 2000 | Difference | Standard Errors Method |
| :---: | :---: | :---: | :---: | :---: |
| 0-2 | 0.4294 | 0.4785 | 0.0491 |  |
|  |  |  | $(0.0074)^{* *}$ | DNP |
|  |  |  | (0.0097)** | NPH |
|  |  |  | (0.0108)** | proposed |
|  |  |  | $(0.0112)^{* *}$ | proposed (no covariance term) |
| 3-6 | 0.4370 | 0.4253 | -0.0117 |  |
|  |  |  | (0.0092) | DNP |
|  |  |  | (0.0125) | NPH |
|  |  |  | $(0.0146)$ | proposed |
|  |  |  | $(0.0150)$ | proposed (no covariance term) |
| 7-11 | 0.6630 | 0.6134 | $-0.0496$ |  |
|  |  |  | $(0.0106)^{* *}$ | DNP |
|  |  |  | $(0.0206)^{* *}$ | NPH |
|  |  |  | $(0.0269)^{*}$ | proposed |
|  |  |  | $(0.0273)^{*}$ | proposed (no covariance term) |
| $12+$ | 0.7264 | 0.6699 | -0.0565 |  |
|  |  |  | $(0.0098)^{* *}$ | DNP |
|  |  |  | $(0.0172)^{* *}$ | NPH |
|  |  |  | (0.0272)** | proposed |
|  |  |  | $(0.0232)^{* *}$ | proposed (no covariance term) |
| total | 0.5355 | 0.5322 | -0.0033 |  |
|  |  |  | (0.0047) | DNP |
|  |  |  | (0.0060) | NPH |
|  |  |  | (0.0079) | proposed |
|  |  |  | (0.0070) | proposed (no covariance term) |

** The estimated difference is significant at the $5 \%$ level

* The estimated difference is significant at the $10 \%$ level
job stability. This result is consistent with the claim that the extent to which the NPH method underestimates the correct variance is correlated with the retention rate. While my method includes an adjustment factor, $f_{k}$, the DNP and NPH methods do not explicitly account for changes in the average weight over time. ${ }^{39}$ This adjustment factor, although conceptually important, is not the driving force behind the results. For the chosen sample, excluding it from my approach would only increase the gap in standard errors, not reduce

[^22]it.

As a result, the DNP and NPH approaches may lead the researcher to falsely reject the null hypothesis of no change in job stability. Calculating t-statistics for males with $12+$ years of tenure illustrates this point. Using either the DNP or NPH methods, one strongly rejects the null hypothesis at the $5 \%$ significance level. In fact, my method suggests that the null hypothesis should not be rejected, even at the $10 \%$ level.

Table 2.4 examines differences in the 4 -year retention across groups for 1996. The results are similar to those found in Table 2.2 and 2.3. The proposed standard errors are consistently larger than the NPH ones, with a larger gap for groups with higher retention rates. In addition, using the correct standard errors can alter the conclusion at the inference stage. With the DNP and NPH methods, one strongly rejects the null hypothesis at the $5 \%$ significance level of no difference in job stability between males and females with 7 to 11 years of initial tenure. Yet my method suggests that the difference is barely significant at the $10 \%$ level.

### 2.11 'Averaged' Retention Rate

In this last section, I define the averaged retention rate, and provide a discussion of its asymptotic properties.

Define $R_{j, j+1}^{a, a^{\prime} ; t, t^{\prime}}$ as the retention rate for a randomly chosen worker from year $j$ or $j+1$ that is aged $a$ to $a^{\prime}$ with $t$ to $t^{\prime}$ months of tenure. Let $y_{i j, j+1}^{a, a^{\prime} ; t, t^{\prime}}$ be a dummy variable which takes the value one if individual i is between $a$ and $a^{\prime}$ years of age and has between $t$ and $t^{\prime}$ months of tenure in period $j$ or $j+1$. Use the shorthand notation $D_{i j, j+1}=y_{i j, j+1}^{a, a^{\prime} ; t, t^{\prime}}$ and $N_{i j, j+1}=y_{i j, j+1}^{a+1, a^{\prime}+1 ; t+12, t^{\prime}+12}$. Therefore, $R_{1,2}^{a, a^{\prime} ; t, t^{\prime}}$ can be written as

$$
R_{1,2}=R_{1,2}^{a, a^{\prime} ; t, t^{\prime}}=\operatorname{Prob}\left(N_{i 2,3}=1 \mid D_{i 1,2}=1\right)
$$

where the underlying population consists of individuals aged 20 to 54 in year 1 and 2 . The constant population assumption assumes one can follow these same individuals one year later, i.e. they will now be aged 21 to 55 in years 2 and 3 .

Table 2.4: 4-year Retention Rates - Group Differentials

| Group Specification |  | $\hat{R}_{1996}^{g_{1}}-\hat{R}_{1996}^{g_{2}}$ | Standard Errors |
| :---: | :---: | :---: | :---: |
| $\underline{g_{1}}$ | $g_{2}$ |  | Method |
| male | female | 0.0442 |  |
|  |  | $(0.0047)^{* *}$ |  |
|  |  | $(0.0061)^{* *}$ | NPH (no covariance term) |
|  |  | $(0.0071)^{* *}$ | , NPH |
|  |  | $(0.0088)^{* *}$ | proposed |
| male; tenure 0-2 | female; tenure 0-2 | 0.0481 |  |
|  |  | $(0.0077)^{* *}$ | DNP |
|  |  | $(0.0101)^{* *}$ | NPH (no covariance term) |
|  |  | $(0.0105)^{* *}$ | NPH |
|  |  | $(0.0123) * *$ | proposed |
| male; tenure 3-6 | female; tenure 3-6 | 0.0152 |  |
|  |  | (0.0094) | DNP |
|  |  | (0.0132) | DNP (no covariance term) |
|  |  | (0.0136) | NPH |
|  |  | (0.0163) | proposed |
| male; tenure 7-11 | female; tenure 7-11 | 0.0439 |  |
|  |  | $(0.0112)^{* *}$ | DNP |
|  |  | $(0.0193) * *$ | NPH (no covariance term) |
|  |  | $(0.0197)^{* *}$ | NPH |
|  |  | $(0.0261)^{*}$ | proposed |
| male; tenure $12+$ | female; tenure $12+$ | 0.0024 |  |
|  |  | (0.0084) | DNP |
|  |  | 0.0166) | NPH (no covariance term) |
|  |  | (0.0172) | NPH |
|  |  | (0.0238) | proposed |

[^23]Proposition 8 Assuming an iid sample for years, and no change in population, then

$$
R_{1,2}=\frac{E\left(N_{i 2,3}\right)}{E\left(D_{k 1,2}\right)}
$$

proof:

$$
\begin{aligned}
R_{1,2} & =\operatorname{Prob}\left(N_{i 2,3}=1 \mid D_{i 1,2}=1\right) \\
& =\frac{\operatorname{Prob}\left(N_{i 2,3}=1\right)}{\operatorname{Prob}\left(D_{i 1,2}=1\right)} \\
& =\frac{\operatorname{Prob}\left(N_{i 2,3}=1\right)}{\operatorname{Prob}\left(D_{k 1,2}=1\right)} \\
& =\frac{E\left(N_{i 2,3}\right)}{E\left(D_{k 1,2}\right)}
\end{aligned}
$$

Let $n_{j}^{D}$ represent the number of individuals that are aged 20 to 54 in year $j$, and $n_{j}^{N}$ represent the number of individuals that are aged 21 to 55 in year $j$. Let $\psi_{1,2}^{1 D}$ and $\psi_{1,2}^{2 D}$ represent the proportion of individuals aged 20 to 54 in the year 1 and 2 population that are in year 1 and 2 , respectively. Further, Let $\psi_{2,3}^{2 N}$ and $\psi_{2,3}^{3 N}$ represent the proportion of individuals aged 21 to 55 in the year 2 and 3 population that are in year 2 and 3 , respectively. Finally, assume individuals denoted by the subscript $k$ are chosen from the sample of 20 to 54 year old, and those denoted by the subscript $i$ from the sample of 21 to 55 year old.

The synthetic cohort estimator is

$$
\hat{R}_{1,2}=\hat{R}_{1,2}^{a, a^{\prime} ; t, t^{\prime}}=\frac{\psi_{2,3}^{2 N}\left(\sum_{i=1}^{n_{2}^{N}} N_{i 2} / n_{2}^{N}\right)+\psi_{2,3}^{3 N}\left(\sum_{i=1}^{n_{3}^{N}} N_{i 3} / n_{3}^{N}\right)}{\psi_{1,2}^{1 D}\left(\sum_{k=1}^{n_{1}^{D}} D_{i 1} / n_{1}^{D}\right)+\psi_{1,2}^{2 D}\left(\sum_{k=1}^{n_{2}^{D}} D_{i 2} / n_{2}^{D}\right)}
$$

Proposition 9 Assuming an iid sample for each year, no change in population, and independence across years, then $\hat{R}_{1,2} \xrightarrow{p} R_{1,2}$
proof: Use the results of Proposition 8

$$
R_{1,2}=\frac{E\left(N_{i 2,3}\right)}{E\left(D_{k 1,2}\right)}
$$

and

$$
\begin{aligned}
& \psi_{2,3}^{2 N}\left(\sum_{i=1}^{n_{2}^{N}} N_{i 2} / n_{2}^{N}\right)+\psi_{2,3}^{3 N}\left(\sum_{i=1}^{n_{3}^{N}} N_{i 3} / n_{3}^{N}\right) \xrightarrow{p} \psi_{2,3}^{2 N} E\left(N_{i 2}\right)+\psi_{2,3}^{3 N} E\left(N_{i 3}\right)=E\left(N_{i 2,3}\right) \\
& \psi_{1,2}^{1 D}\left(\sum_{k=1}^{n_{1}^{D}} D_{k 1} / n_{1}^{D}\right)+\psi_{1,2}^{2 D}\left(\sum_{k=1}^{n_{2}^{D}} D_{k 2} / n_{2}^{D}\right) \xrightarrow{p} \psi_{1,2}^{1 D} E\left(D_{k 1}\right)+\psi_{1,2}^{2 D} E\left(D_{k 2}\right)=E\left(N_{k 1,2}\right)
\end{aligned}
$$

one can conclude

$$
\frac{\psi_{2}^{N} \sum_{i=1}^{n_{2}^{N}} N_{i 2} / n_{2}^{N}+\psi_{3}^{N} \sum_{i=1}^{n_{3}^{N}} N_{i 3} / n_{3}^{N}}{\psi_{1}^{D} \sum_{k=1}^{n_{1}^{D}} D_{k 1} / n_{1}^{D}+\psi_{2}^{D} \sum_{k=1}^{n_{2}^{D}} D_{k 2} / n_{2}^{D}} \xrightarrow{p} \operatorname{Prob}\left(N_{i 2,3}=1 \mid D_{i 1,2}=1\right)
$$

Proposition 10 provides the asymptotic properties of the averaged retention rate.

Proposition 10 Assuming an iid sample for each year, no change in population, independence across years, and $\lim _{n_{1}^{D}, n_{k}^{r} \rightarrow \infty} \frac{n_{1}^{D}}{n_{k}^{r}}=1$, for $k=1,2$ and $r=D, N$ then $\sqrt{n_{1}^{D}}\left(\hat{R}_{1,2}-R_{1,2}\right) \xrightarrow{d} N(0, V)$ where $V$ is

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(D_{i 2}\right)+\phi_{3}^{2} V\left(N_{i 2}\right)+\phi_{4}^{2} V\left(N_{i 3}\right)-2 \phi_{2} \phi_{3} \mu C\left(N_{i 2}, D_{i 2}\right) \tag{2.21}
\end{equation*}
$$

with

$$
\begin{aligned}
\phi_{1} & =\psi_{1,2}^{1 D} \frac{N_{i 2,3}}{E\left(D_{k 1,2}\right)^{2}} \\
\phi_{2} & =\psi_{1,2}^{2 D} \frac{N_{i 2,3}}{E\left(D_{k 1,2}\right)^{2}} \\
\phi_{3} & =\psi_{2,3}^{2 N} \frac{1}{E\left(D_{k 1,2}\right)} \\
\phi_{4} & =\psi_{2,3}^{3 N} \frac{1}{E\left(D_{k 1,2}\right)}
\end{aligned}
$$

and $\mu$ is the probability that a randomly chosen person from the population of individuals aged 20 to 54 in year 2, will be 21 to 54.
proof: For ease of notation let $\hat{N}_{2,3}=\psi_{2,3}^{2 N} \hat{N}_{i 2}+\psi_{2,3}^{3 N} \hat{N}_{i 3}, N_{2,3}=\psi_{2,3}^{2 N} N_{i 2}+\psi_{2,3}^{3 N} N_{i 3}$,
and define $\hat{D}_{1,2}, D_{1,2}$ in a similar fashion.

$$
\begin{aligned}
& \sqrt{n_{1}^{D}}\left(\hat{R}_{1,2}^{a, a^{\prime} ; t, t^{\prime}}-\hat{R}_{1,2}^{a, a^{\prime} ; t, t^{\prime}}\right) \\
& =\sqrt{n_{1}^{D}}\left(\frac{\hat{N}_{2,3}}{\hat{D}_{1,2}}-\frac{N_{2,3}}{D_{1,2}}\right) \\
& =\sqrt{n_{1}^{D}} \frac{\left(\hat{N}_{2,3}-N_{2,3}\right) D_{1,2}-\left(\hat{D}_{1,2}-D_{1,2}\right) N_{2,3}}{D_{1,2} \hat{D}_{1,2}} \\
& +\sqrt{n_{1}^{D}} \frac{\left(\hat{N}_{2,3}-N_{2,3}\right) D_{1,2}-\left(\hat{D}_{1,2}-D_{1,2}\right) N_{2,3}}{D_{1,2}^{2}}+o_{p}(1) \\
& =-\phi_{1} \sqrt{n_{1}^{D}}\left(\hat{D}_{1}-D_{1}\right)-\phi_{2} \sqrt{n_{2}^{D}}\left(\hat{D}_{2}-D_{2}\right)+\phi_{3} \sqrt{n_{2}^{N}}\left(\hat{N}_{2}-N_{2}\right)+\phi_{4} \sqrt{n_{3}^{N}}\left(\hat{N}_{3}-N_{3}\right)+o_{p}(1) \\
& \xrightarrow{d} N\left(0, \phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(D_{i 2}\right)+\phi_{3}^{2} V\left(N_{i 2}\right)+\phi_{4}^{2} V\left(N_{i 3}\right)-2 \phi_{2} \phi_{3} \mu C\left(D_{i 2}, N_{i 2}\right)\right)
\end{aligned}
$$

The retention rate that accounts for non-randomness and a change, in population is

$$
\tilde{R}_{1,2}=\frac{\sum_{j=2}^{3} \sum_{i=1}^{n_{j}^{N}} b w_{i j} N_{i j}}{\sum_{j=1}^{2} \sum_{k=1}^{n_{j}^{D}} b w_{k j} D_{k j}}
$$

Alternatively, the estimator can be rewritten as

$$
\tilde{R}_{1,2}=\operatorname{adj} j_{1,2} \frac{\hat{\psi}_{2,3}^{2 N}\left[\sum_{i=1}^{n_{2}^{N}} n w_{i 2} N_{i 2} / n_{2}^{N}\right]+\hat{\psi}_{2,3}^{3 N}\left[\sum_{i=1}^{n_{3}^{N}} n w_{i 3} N_{i 3} / n_{3}^{N}\right]}{\hat{\psi}_{1,2}^{1 D}\left[\sum_{k=1}^{n_{1}^{D}} n w_{k 1} D_{k 1} / n_{1}^{D}\right]+\hat{\psi}_{1,2}^{2 D}\left[\sum_{k=1}^{n_{2}^{D}} n w_{k 2} N_{k 2} / n_{2}^{D}\right]}
$$

where

$$
\begin{gathered}
\hat{\psi}_{1,2}^{1 D}=\frac{\sum_{k=1}^{n_{1}^{D}} b w_{k 1}}{\sum_{k=1}^{n_{1}^{D}} b w_{k 1}+\sum_{k=1}^{n_{2}^{D}} b w_{k 2}}, \quad \hat{\psi}_{1,2}^{2 D}=\frac{\sum_{k=1}^{n_{2}^{D}} b w_{k 2}}{\sum_{k=1}^{n_{1}^{D}} b w_{k 1}+\sum_{k=1}^{n_{2}^{D}} b w_{k 2}} \\
\hat{\psi}_{2,3}^{2 N}=\frac{\sum_{i=1}^{n_{2}^{N}} b w_{i 2}}{\sum_{i=1}^{n_{2}^{N}} b w_{i 2}+\sum_{i=1}^{n_{3}^{N}} b w_{i 3}}, \quad \hat{\psi}_{2,3}^{3 N}=\frac{\sum_{i=1}^{n_{3}^{N}} b w_{i 3}}{\sum_{i=1}^{n_{2}^{N}} b w_{i 2}+\sum_{i=1}^{n_{3}^{N}} b w_{i 3}}
\end{gathered}
$$

and

$$
a d j_{1,2}=\frac{\sum_{j=2}^{3} \sum_{i=1}^{n_{j}^{N}} b w_{i j}}{\sum_{j=1}^{2} \sum_{k=1}^{n_{k}^{D}} b w_{k j}}
$$

This estimator shares the same asymptotic properties as $\hat{R}_{1,2}$, but the variance estimator
needs adjustment to account for the population change, i.e. not following a population cohort. Treating $a d j_{1,2}$ as a constant and $\hat{\psi}_{1,2}^{1 D}, \hat{\psi}_{1,2}^{2 D}, \hat{\psi}_{2,3}^{2 N}$ and $\hat{\psi}_{2,3}^{3 N}$ as true measures of the population proportions, the variance estimator is the (weighted) sample analogs of (2.21), pre-multiplied by $\left(a d j_{j}\right)^{2}$.

### 2.12 Conclusion

The main contribution of this Chapter is to provide a method for calculating correct tstatistics with repeated cross sectional data. Using this method one can properly identify breaks in job stability patterns. I show that the retention rate can be consistently estimated using cross sectional data, and provide a clear and intuitive discussion of identifying assumptions. I also propose a method for consistently estimating standard errors, and show, both theoretically and empirically (using CPS data) that existing methods for estimating standard errors have a downward bias-a bias that may lead to a spurious identification of breaks in job stability patterns.

## Chapter 3

## Job Stability Patterns

### 3.1 Introduction

Existing North American literature has arrived at a consensus that job stability has not declined in the 1980s and 1990s, as was previously believed. There is also some evidence of changes over time in stability relative to race, education and gender. ${ }^{1}$ Yet data limitations, and the lack of precision in existing tools, make it difficult to differentiate between cyclical and secular changes, thereby hampering the forward progress of this literature. ${ }^{2}$ Without a clear identification of breaks in job stability patterns; one cannot attempt to explore their causes.

In this paper, I document changing job stability patterns in Canada over the 19772004 period using the master Labour force Survey (LFS) files. The LFS is the only North American data source to have consistent and detailed tenure data on a regular basis, and more importantly, over an extended period of time. Using this data, and tools developed in Chapter 2, I find that what was previously seen as cyclical change is actually a secular 1 increase in job stability. The main finding reveals that job stability has increased since the early 1990s-at the aggregate level, and especially for women and workers with less than one year of initial tenure. ${ }^{3}$

[^24]In this paper I also document job stability patterns over a wide range of sub-populations, some of which have yet to be explored. Regions and industry are examples of dimensions which have neither been identified nor scrutinized. Providing such a thorough exposition of the patterns helps determine the extent of the changes experienced since the early 1990 s, and as I argue in Chapter 4, helps identify their sources.

A final contribution of this paper is the comparison of Canadian and American job stability patterns. Heisz (2005) has systematically compared patterns across the two countries, but only up to 1995. By extending the comparison period to include the late 1990s and and early 2000 s, one can see whether the changes experienced in Canada since the early 1990 s are also present south of the border.

The structure of this chapter is as follows: Section 3.2 describes the main data set used in the empirical analysis. Section 3.3 provides a characterization of job stability patterns for different Canadian sub-populations, with a focus on ageing and gender as potential sources of the new job stability patterns. Section 3.4 carries out a Canada-U.S. comparison.

### 3.2 LFS Data

The main data used in this Chapter is from the master Labour Force Survey (LFS) files. ${ }^{4}$ The LFS is a large monthly household survey, approximately 54,000 households per month, with a focus on gathering information about the labour market activities of Canadians. The unique strength of the LFS is the quality and consistency of tenure data. As part of the regular LFS questionnaire, dating back to 1976, employed respondents are asked, "When did he/she start working for [name of employer]". The LFS interviewer manual describes in detail how to count tenure. For casual work, one week away breaks up a tenure spell. For seasonal work, the tenure spell breaks when the worker does not work in the off-season. In all cases there is a break if they do work but at a job with a different employer. Therefore,

[^25]the tenure question is meant to measure the most recent period of uninterrupted work at one employer. ${ }^{5}$

The sample contains all individuals between the ages of 20 and 55 in the incoming rotation group over the 1977-2004 period. The LFS follows a rotating panel design, where a household remains in the sample for six consecutive months and every month one sixth of the sample is replaced. By restricting attention to the incoming group one can ensure a random sample. The upper age limit accounts for the falling retirement age. By excluding those above the age of 55 , one can focus on quits and layoffs, and not have the results tainted by voluntary retirement. Because only a small fraction of 15-19 year olds work, with the majority still in school, the sample was started at the age of 20 . Although data is available as of January 1976, the 1976 files are much smaller than subsequent years; 35.7\% smaller than in 1977, and $32.5 \%$ smaller than the 1977-1981 average. ${ }^{6}$ The start year is important in its role of determining the group proportions for the decomposition exercises. Considering the 29 years of available data, the benefits of starting the sample in 1977, outweigh the costs of discarding some information.

The synthetic cohort approach, as presented in Chapter 2, requires the inclusion of nonworkers. ${ }^{7}$ However, this does not limit the range of retention rates that can be estimated.

[^26]One can still construct a retention rate that conditions on a narrow set of characteristics. ${ }^{8}$ In this Chapter, the minimum conditioning characteristic common to all retention rates include: all currently employed individuals between the ages of 20 and 54 , except full-time students, the self-employed and those in the military. Full-time students, as well as young adults that worked during the summer months but intended to go back to school in the fall, are not part of this study because working was not their main activity. The self-employed and those working in the military were also conditioned out because the process determining their job tenure spell is very different from (non-military) paid employees.

### 3.3 Canadian Job Stability Patterns

This section documents the Canadian job stability patterns for the 1977-2004 period using a retention rate approach. A detailed explanation of the retention rate approach and the estimator used can be found in Chapter 2.

Figure 3.1 shows the overall one-year retention rate for 1977-2003. ${ }^{9}$ In the early part of the sample, the overall retention rate was counter-cyclical, averaging $78.7 \%$ over the 19771989 period. But by the early 1990s the pattern had changed. The retention rate increased substantially in the 1990s, well above $80 \%$, and stayed fairly constant at these historically high levels for the duration of the 1990s and into the 2000s. As illustrated in Figure 3.1, the narrow width of the 95 percent confidence interval clearly indicates a break in the aggregate pattern in the 1990s. ${ }^{10}$

[^27]

Figure 3.2 shows differences in retention rates between men and women. Prior to the 1990s, the male retention rate had exceeded the $80 \%$ mark on more than one occasion, but never remained above this threshold for long. For females it was a different story. The retention rates of the 1990s and 2000s are well above anything achieved prior to this period. Retention rates were also calculated by birth cohorts, to better understand the transition to the new job stability patterns. To ensure more accurate estimates, 2 -year cohorts were used (e.g., the retention rates for workers born in 1957 and 1958 were calculated for 1977, $1979, \ldots, 2001$ ). For the gender breakdown, there was one clear pattern. The increase in stability for women in the 1990s was not restricted to newer cohorts, but was present across birth cohorts. It is also important to note that for both sexes, retention rates stabilized in the latter part of the sample. This shows that an important part of the new patterns is not gender driven.


Figure 3.3 clearly indicates that the positive male-female retention rate differential in the early part of the sample was both economically and statistically significant. In the latter part of the sample, however, the positive gap disappeared and in many cases turned negative (although in most cases not statistically significant).


Figure 3.4 shows important differences across age groups-revealing that older workers have consistently higher job stability than younger workers. This figure also points to a possible concave relationship between age and stability. To further explore this possibility, I estimated retention rates for 5 -year age intervals. The results support this claim. Job stability increases with age, but at a decreasing rate, eventually peaking when the worker reaches his late forties. ${ }^{11}$ The gains in stability are still economically significant as the worker goes through his 30 s; there remains a consistent 5 percentage point gap between 30 34 and 35-39 year olds throughout the sample period. These findings indicate that ageing of the workforce may have an important role to play in explaining the overall increase in stability-a point that is further explored in Chapter 4.

[^28]

Figure 3.5 provides an education breakdown. ${ }^{12}$ Starting in 1990, the LFS introduced some important modifications to its education questions. The focus changed from measuring years of education to measuring education attainment. As a result, the construction of time consistent education groupings is problematic. At best, the problem can only be attenuated. ${ }^{13}$ The three education attainment categories for Figure 3.5 are high school or less, post secondary diploma or certificate, and university degree (bachelor or more). The categories are self-explanatory, with one exception. Respondents with some post secondary, but who did not complete their program, were included in the high school or less group. ${ }^{14}$ This graph indicates individuals with high school or less have systematically lower retention rates. ${ }^{15}$ Although caution must be taken when comparing trends in the pre- and post-1990

[^29]periods, it also appears that the counter-cyclical patterns for both high and low levels of education, present in the first half of the sample, have disappeared. Interestingly, the results for the low education group are robust to the exclusion of post secondary dropouts. This finding supports the perception among labour economists that post secondary dropouts have more in common with high school graduates than those who successfully completed a post secondary program.


Figures 3.6 through 3.10 show that the new job stability patters are broadly based. Irrespective of whether one carries a regional or industry breakdown, or even compares private and public sectors, retention rates have tended to stabilize at high levels in the 1990s and 2000s. Some interesting cross sector differences do surface, however. Figure 3.6 shows a significant decline prior to recessions for the goods sector. ${ }^{16}$ A breakdown of the goods sector, see Figure 3.7, reveals that the more seasonal the industry, the more

[^30]unstable the employment relationship. Retention rates for the agriculture, forestry, fishing and hunting sector are, on average, 20 percentage points lower than in manufacturing. This finding is reflected in Figure 3.9, with Atlantic Canada and its greater dependence on seasonal employment consistently having lower retention rates as compared to Ontario and Quebec. Figure 3.8 separates the private from the public sector, showing that the public sector tends to be more stable with a difference of approximately 10 percentage points. This gap may be explained by the higher unionization rate in the public sector. Figures 3.9 and 3.10 demonstrate regional differences. Quebec and Ontario mirror each other quite closely as they continually exceed retention rates from other regions. The retention rates for the Prairies show a marked increase since the 1970s.




Figure 3.11 breaks down the retention rate into union and non-union sectors, ${ }^{17}$ but only for the 1997-2004 period. Prior to 1997, a union question was not part of the regular LFS questionnaire. As a result, one cannot comment on potential long term changes. But Figure 3.11 does clearly indicate that jobs were more stable in the union sector, with the gap ranging from 12-14 percentage points over the 1997-2004 period. This result is consistent with the belief that unions safeguard jobs. The presence of a union could also affect quit behaviour. A union provides a collective voice for the workers; it provides a way for workers to voice their discontent, and as such, may reduce voluntary turnover. ${ }^{18}$


A major finding of this Chapter is that conditioning on initial tenure dramatically alters the job stability patterns. Figure 3.12 illustrates that for much of the first part of the sample, the retention rate for workers with less than one year of initial tenure hovered in the 44-49\% range, but in the 1990s this changed dramatically. A steep 10 percentage point increase

[^31]began in 1992. The confidence interval in Figure 3.12 points to a statistically significant change.


For those with initial tenure of one year or more, see Figure 3.13, there is no break in the pattern, although the most recent decline is not as steep as prior ones. . Figure 3.14 shows that the cyclical patterns are present in medium term jobs ( 3 to 7 years of initial tenure), but there is no clear trend for longer term jobs.


Figure 3.14: One Year Retention Rates


The results of Figures 3.12 through 3.14 clearly show that an important part of the changes are concentrated at the low end of the tenure distribution. Given the richness of the LFS data, one can take the identification process one step further by constructing retention rates that condition on months - and not just years - of initial tenure. ${ }^{19}$ Figure 3.15 shows that the increases in stability are more dramatic in the first few months of initial tenure. From 1993 to 2000, retention rates for workers with initial tenure of two months or less increased by 20 percentage points. This figure also indicates that once one conditions on at least 6 months of initial tenure, the increases have all but dissipated.


An interesting point to note from Figure 3.15 is that retention rates initially decrease with tenure. In all but one year, the one-ýear retention rate for workers with one to two months of initial tenure were systematically higher than the retention rates for workers with three to five months of initial tenure. This is consistent with the findings of researchers (e.g. Farber (1994)) that have estimated monthly retention rates. ${ }^{20}$ These studies find that the

[^32]retention rate bottoms at approximately 3 months, and then begin a slow ascent. Farber (1999) argues that the initial decrease and subsequent increase is consistent with a Jovanovic (1979) type match quality model. The worker is less willing to leave the job very early due to the option value of the match-there is a possibility that the match may turn out to be good.

In Figures 3.16 and 3.17, I break down the male and female retention rates by initial tenure. Figure 3.16 shows that the increase in stability for newer jobs is not gender driven. Both males and females experienced significant increases in stability starting in the early 1990s. Surprisingly, Figure 3.16 also shows that the female retention rate for newer jobs is consistently higher that its male counterpart-even in the late 1970s and early 1980s. The decline in the male-female retention rate differential observed in Figure 3.3 must have occurred in longer tenured jobs-a conjecture confirmed in Figure 3.17.

one minus the hazard, this does not materially alter the argument.


Additional breakdowns for those with less than one year of initial tenure (see Figures 3.18 through 3.24 ) reveal that the dramatic increase in stability for newer jobs is broadly based-as was the case for the overall retention rate. There are two important differences, the first of which has already been mentioned. Gender appears to play a role in explaining the new aggregate patterns, but not for new jobs. A similar story appears to hold for age. Figure 3.18 shows that the large increase in stability of newer jobs is present across all age groups. The importance of gender and ageing will be quantified in Chapter 4 using decomposition techniques.

Figure 3.18: One Year Retention Rates



Figure 3.22: One Year Retention Rates



$$
\begin{aligned}
& - \text { public sector } \\
& \\
& \text { private sector }
\end{aligned}
$$




### 3.3.1 Controlling for the Business Cycle

By comparing retention rates at similar stages of the business cycle, one can put in perspective the importance of changes in job stability. This approach takes into account systematic changes in firm or employee behavior that are linked to the business cycle. For example, employees may be more reticent to leave jobs in recessions when outside job prospects are more limited. It also addresses self-selection problems associated with systematic changes in the composition of the workforce. By comparing two periods one can also formally test for changes in job stability. Tables 3.1, 3.2 and 3.3 compare the 1987-1989 and 1998-2000 periods, both strong expansionary periods, which precede and follow dramatic changes in job stability. A weighted average of two consecutive years' retention rates was used to minimize the sensitivity of results to the choice of start and end years. ${ }^{21}$

Table 3.1 summarizes the main findings such as the increase in aggregate job stability, and in particular, the marked increase in job stability for women and workers with less than

[^33]one year of initial tenure. For the latter group, the change is particularly striking. After having controlled for the position of the business cycle, there still remains a 10.1 percentage point rise, representing a $22.7 \%$ increase from a decade ago.

Table 3.1: Comparison Across Business Cycles

| Specification | $R_{1987,1988}$ | $R_{1998,1999}$ | $\left(R_{1998,1999}-R_{1987,1988}\right)$ |
| :--- | :---: | :---: | :---: |
| Overall | 0.777 | 0.821 | 0.045 |
|  | $(0.002)$ | $(0.002)$ | $(0.002)$ |
| Male | 0.786 | 0.819 | 0.033 |
|  | $(0.003)$ | $(0.003)$ | $(0.004)$ |
| Female | 0.767 | 0.822 | 0.055 |
|  | $(0.003)$ | $(0.003)$ | $(0.004)$ |
| Tenure < 1 year | 0.444 | 0.545 | 0.101 |
|  | $(0.004)$ | $(0.005)$ | $(0.006)$ |

Table 3.2 shows that the new job stability patterns are broadly based. Irrespective of whether one carries out an industry breakdown, ${ }^{22}$ or even compares private and public sectors, retention rates have tended to stabilize at high levels in the 1990s and 2000s. The one exception is the Agriculture, Forestry, Fishing and Hunting sector. Given the strong seasonal and environmental dependence of this sector, an insignificant result is not surprising. Table 3.2 also shows that the new job stability patterns are present across the country. These changes are not only economically significant, but also statistically significant. In all but one case, one can strongly reject the null hypothesis of no change at the $5 \%$ significance level.

[^34]Table 3.2: Comparison Across Business Cycles, by Sector and Region

| Specification | $R_{1987,1988}$ | $R_{1998,1999}$ | $\left(R_{1998,1999}-R_{1987,1988}\right)$ |
| :--- | :---: | :---: | :---: |
| Sector |  |  |  |
| Goods | 0.784 | 0.826 | 0.042 |
|  | $(0.004)$ | $(0.004)$ | $(0.006)$ |
| Services | 0.774 | 0.819 | 0.045 |
|  | $(0.002)$ | $(0.002)$ | $(0.003)$ |
| Agriculture, Forestry, |  |  |  |
| Fishing and Hunting | 0.641 | 0.639 | -0.001 |
|  | $(0.016)$ | $(0.016)$ | $(0.023)$ |
| Mining, Oil and Gas Extraction, |  |  |  |
| Utilities and Construction | 0.740 | 0.770 | 0.029 |
|  | $(0.007)$ | $(0.009)$ | $(0.011)$ |
| Manufacturing | 0.816 | 0.862 | 0.046 |
|  | $(0.005)$ | $(0.005)$ | $(0.007)$ |
| Public | 0.868 | 0.922 | 0.054 |
|  | $(0.004)$ | $(0.005)$ | $(0.006)$ |
| Private | 0.745 | 0.790 | 0.045 |
|  | $(0.002)$ | $(0.002)$ | $(0.003)$ |
| Region |  |  |  |
| Atlantic | 0.752 | 0.770 | 0.018 |
|  | $(0.008)$ | $(0.008)$ | $(0.011)$ |
| Quebec | 0.775 | 0.836 | 0.061 |
|  | $(0.004)$ | $(0.004)$ | $(0.006)$ |
| Ontario | 0.795 | 0.832 | 0.037 |
|  | $(0.003)$ | $(0.003)$ | $(0.005)$ |
| Prairies | 0.749 | 0.798 | 0.049 |
| British Columbia | $(0.005)$ | $(0.005)$ | $(0.007)$ |
|  | 0.775 | 0.817 | 0.042 |

Table 3.3 shows that the new pattern job stability patterns for workers with initial tenure of less than one year are broadly based. In all but two cases (Agriculture, Forestry, Fishing and Hunting; and Mining Oil and Gas Extraction, Utilities and Construction being the exceptions), the increases are both economically and statistically significant.

Table 3.3: Comparison Across Business Cycles - Conditioning on Initial Tenure Less than

| 1 Year | $R_{1987,1988}$ | $R_{1998,1999}$ | $\left(R_{1998,1999}-R_{1987,1988}\right)$ |
| :--- | :---: | :---: | :---: |
| Specification | 0.402 | 0.506 | 0.104 |
| Sector | $(0.007)$ | $(0.010)$ | $(0.012)$ |
| Goods | 0.462 | 0.559 | 0.098 |
|  | $(0.005)$ | $(0.006)$ | $(0.008)$ |
| Services |  |  |  |
|  | 0.184 | 0.230 | 0.046 |
| Agriculture, Forestry, | $(0.015)$ | $(0.019)$ | $(0.024)$ |
| Fishing and Hunting |  |  |  |
|  |  | 0.344 | -0.001 |
| Mining, Oil and Gas Extraction, | 0.345 | $(0.018)$ |  |
| Utilities and Construction | $(0.011)$ | $(0.014)$ | 0.163 |
|  | 0.475 | 0.638 | $(0.018)$ |
| Manufacturing | $(0.011)$ | $(0.0148)$ | 0.109 |
|  | 0.520 | 0.628 | $(0.020)$ |
| Public | $(0.012)$ | $(0.016)$ | 0.102 |
|  | 0.430 | 0.532 | $(0.007)$ |
| Private | $(0.004)$ | $(0.005)$ |  |
|  |  |  | 0.048 |
| Region | 0.338 | 0.386 | $(0.018)$ |
| Atlantic | $(0.012)$ | $(0.014)$ | 0.125 |
|  | 0.415 | 0.539 | $(0.013)$ |
| Quebec | $(0.008)$ | $(0.011)$ | 0.077 |
|  | 0.513 | 0.590 | $(0.011)$ |
| Ontario | $(0.007)$ | $(0.009)$ | 0.108 |
|  | 0.415 | 0.523 | $(0.015)$ |
| Prairies | $(0.010)$ | $(0.012)$ | 0.167 |
| British Columbia | 0.396 | 0.564 | $(0.019)$ |
|  | $(0.011)$ | $(0.015)$ |  |

### 3.4 Canada - U.S. Comparison

This section compares Canadian and American job stability patterns. These two economies have experienced many similar changes over the last half-century. These include: a large post-war baby-boom; a significant increase in the participation rate of women; and until recently, similar timings of business cycles; to name just a few. Only Heisz (2005) has systematically compared the long term changes in job stability for the two countries. Examining the 1983 to 1995 period, he found that neither country experienced a period-long decline in job stability, but both saw an increase in stability for women relative to men.

In this Chapter I extend the comparison period to include the late 1990s and early 2000 s. Limiting the Canadian data to no later than 1995 leads to a very different interpretation of the Canadian patterns. The large increase in Canadian job stability experienced in the early 1990s would have been attributed to the 1990-1991 recession and the subsequent slow recovery. Job stability in Canada had tended to be counter-cyclical, increasing in recessions only to fall back in periods of expansion, and this would have appeared as a continuation of the same pattern. Only with the inclusion of data for the late 1990s and early 2000s did it become clear that new Canadian patterns had emerged. In addition, this comparison will help determine whether there are common trends irrespective of age restrictions. For this study, the focus is on prime age workers, i.e. 20 to 55 years of age, while in the literature no age restriction is typically imposed.

For the American comparison, I used the public access CPS files. The LFS and CPS are similar in size and scope, making for a natural comparison. Both are large scale data sets, approximately 50,000 household interviews per month, and both gather monthly labour market information on their respective populations. For this Chapter, the U.S. sample includes all individuals aged 20 to 55 . Due to poor tenure data, the minimum conditioning characteristics previously used could not be replicated for the United States. As a result, a modified set of minimum conditioning characteristics were imposed to both countries to ensure compatibility. These included all workers aged 20 to 51 , except for the unincorporated self-employed. ${ }^{23}$ CPS tenure data is only available at irregular intervals. A tenure question is not part of the monthly CPS questionnaire, but is only included in select supplements. To ensure consistency in the tenure question across time, only the January 1983, January 1987, January 1991, February 1996, February 2000 and January 2004 CPS Supplements were used. ${ }^{24}$ Finally, only the January LFS files (all rotation groups) were used to limit

[^35]seasonal differences.
Four-year retention rates were estimated for Canada and the United States. By measuring job stability over a four year interval, a consistent comparison could be made across time and country. For the United States, this approach generates only four point estimates: $\hat{R} 4_{1983}, \hat{R} 4_{1987}, \hat{R} 4_{1996}$ and $\hat{R}_{2000} .{ }^{25}$ For Canada, much of the data richness is set aside. The four-year estimator does not make a distinction between job separations occurring in the first or fourth year. It only identifies the probability an employment relationship will exist four years later.

Figures 3.25 through 3.28 compare four-retention rates for the two countries. For Canada, the general patterns found in the one-year rates still appear. The male, female and overall four-year rates initially exhibit cyclical patterns but stabilize at high levels in the 1990 s and 2000 s . The patterns for the 1983-1991 data are similar to what was found in the rest of the literature - no large scale decline in job stability. ${ }^{26}$ There is some evidence of an increase in female job stability compared to males, but not as pronounced as in Canada.

[^36]



The four-year American retention rates are consistently lower than in Canada. ${ }^{27}$ Differences in union coverage may be an intriguing potential cause. Historically, Canada has had higher union density, and as previously discussed, union covered jobs are more stable.

Table 3.4 shows how much of the gap between U.S. and Canadian 4-year retention rates in 2000 can be explained due to differences in union coverage. Data limitations restrict the comparison to the year 2000 only. ${ }^{28}$ A compositional effect of $100 \%$ would mean that differences in union coverage can fully account for the retention rate differential between the two countries. ${ }^{29}$

Table 3.4: Canada-U.S. Differential in the year 2000

| Specification | Compositional <br> effect |
| :--- | :---: |
|  | $(\%)$ |
| Overall | 84.1 |
| Male | 114.3 |
| Female | 70.1 |
| Three years of tenure or less | 188.4 |

The results in Table 3.4 indicate that union coverage matters. A large part of the aggregate retention rate gap can be attributed to higher union coverage in Canada. Table 3.4 also shows that union coverage matters more for men than females. However, since the changes in union coverage over-explains the gap for males and particularly for workers with three or less years of tenure, this does indicate other factors are at work.

The limited number of American observations, all at different stages in the business cycle, make it all but impossible to interpolate between data points and draw precise conclusions. Focussing on the same data points for Canada would have led us to believe that the rates of the 1990 s and 2000 s were no different than in the early part of the sample. However, there is also a clear indication of a change at low initial tenure in the United States. The extension of the data to the mid 2000s indicates that the increase in stability for newer jobs

[^37]may be part of a secular change, and not just part of a cyclical pattern. Interestingly, the United States has seen an erosion of the "Employment at Will" doctrine over this same period (see Jones and Kuhn (1995) and Autor (2003) for more details), a change that will lead to higher firing costs. In a world where jobs are both an experience and inspection good, the increase in stability of newer jobs is consistent with an increase in job match quality brought about by those higher firing costs. The infrequency of the CPS tenure data prevents the exploration of this hypothesis. As a result, I plan to use the the SIPP (from 1990 onwards ${ }^{30}$ ) to explore this issue in future research.

### 3.5 Conclusion

The main contribution of this Chapter is the documentation of striking new job stability patterns in Canada. Using a rich source of tenure data and tools developed for this Chapter, I differentiate between secular and cyclical changes, and find that aggregate job stability has increased since the early 1990s. The overall retention rate was cyclical, averaging $78.7 \%$ over the 1977-1989 period. But in the early 1990s it increased substantially to well above $80 \%$, and stayed fairly constant at these historically high levels into the 2000s. These changes are economy-wide, with a more significant increase for women and for workers with initial tenure of less than one year.

Results also indicate that ageing of the workforce and gender differences may play an important role in explaining the new overall patterns, but these differences have no impact upon newer jobs.

Finally, the limited number of American observations, all at different stages of the business cycle, make it difficult to draw broadly based conclusions when compared with Canada. However, with the extension of the data to the mid 2000s, there is some indication of change at low initial tenure in the United States.

[^38]
## Chapter 4

## Explaining the New Canadian Job

## Stability Patterns

### 4.1 Introduction

In this chapter, I explore the sources of the new Canadian job stability patterns documented in Chapter 2. Although the main focus is the aggregate pattern, I also explore the causes of the dramatic increase in stability experienced by workers with less than one year of initial tenure. The detailed accounting of possible explanations includes: decline of unions, changes in industry structure, decline in seasonal jobs, increasing education levels, increased labour force attachment of women, ageing of the workforce, and changing match quality.

I use an Oaxaca-Blinder decomposition approach to quantify the importance of compositional changes on the overall retention rate. Said differently, I create counterfactual retention rates where the composition of the workforce is kept constant over time, and compare them to actual rates. ${ }^{1}$ I find that ageing of the workforce can explain almost half of the increase in overall stability from the late 1980 s to the late 1990s. I also find that the

[^39]increased labour force attachment of women and the increased educational attainment of workers also matter. For the latter, the declining importance of workers with low levels of education is the key, and not the rising proportion of workers with university degrees.

This Chapter shows that the application of standard decomposition techniques to subpopulation patterns will not address the counterfactuals of interest. Specifically, it will not control for compositional changes of the workforce. As part of my thesis, I provide an extension that can account for these compositional changes in more narrowly defined rates. I find that ageing cannot explain the large increase in stability of newer jobs. I do find that the rising educational levels of workers can account for part of this increase, but much remains unexplained.

I use a match quality framework where a job is both an inspection and experience good to more thoroughly explore the changes in stability along tenure. Within the context of this model, I show that the stricter EI eligibility requirements introduced in the early 1990s had an affect on match formation. Job seekers are now more selective in their acceptance of job offers, leading to fewer low quality matches being formed.

The structure of the chapter is as follows: Section 4.2 presents the decomposition tools; Section 4.3 provides a detailed accounting of possible structural explanations for the changes at the aggregate level; Section 4.4 carefully examines changes at low levels starting with a decomposition approach, and then proceeds to a match quality search framework; Section 4.5 concludes.

### 4.2 Retention Rate Tools

In this section I present the tools used to explore the causes of new job stability patterns documented in Chapter 2. I start, in Section 4.2.1, with a formal presentation of existing tools, and show how they can help determine the cause of aggregate patterns. In Section 4.3.2, I show how these same tools, when applied to more narrowly defined retention rates, will not provide the counterfactuals of interest. Finally, I propose a refinement that can correctly identify the source of change at the subpopulation level, and in particular, help
identify the source of the important increase in stability of newer jobs.

### 4.2.1 Aggregate Patterns

Due to its straightforward mathematical representation, i.e. a conditional probability of a binary variable, the retention rate can be rewritten to identify the source of the new aggregate job stability patterns. The following known proposition illustrate this point. ${ }^{2}$

Proposition 11 The retention rate, $R_{j}$, can be expressed as a linear combination of retention rates for any G sub-groups

$$
\begin{equation*}
R_{j}=\sum_{k=1}^{G} \gamma_{j}^{g_{k}} R_{j}^{a, \alpha^{\prime} ; t, t^{\prime} ; g_{k}} \tag{4.1}
\end{equation*}
$$

as long as the $G$ groups are mutually exclusive and exhaust the conditioned population (i.e. each individual in year $j$ is between a and $a^{\prime}$ years of age and has between $t$ and $t^{\prime}$ months of tenure), with $\gamma_{j}^{g_{k}}$ representing the proportion of the conditioned population in year $j$ that is in group $g_{k}$.
proof: Use the notation shorthand $N_{i j}^{g_{k}}=y_{i j}^{a+1, a^{\prime}+1 ; t+12, t^{\prime}+12 ; g_{k}}$, and $D_{i j}^{g_{k}}=y_{i j}^{a, a^{\prime} ; ;, t^{\prime} ; g_{k}}$

$$
\begin{aligned}
R_{j} & =\operatorname{Prob}\left(N_{i j+1}=1 \mid D_{i 1}=1\right) \\
& =\frac{\operatorname{Prob}\left(N_{i j+1}=1\right)}{\operatorname{Prob}\left(D_{i j}=1\right)} \\
& =\sum_{k=1}^{G}\left(\frac{\operatorname{Prob}\left(D_{i j}^{g_{k}}=1\right)}{\operatorname{Prob}\left(D_{i j}=1\right)}\right)\left(\frac{\operatorname{Prob}\left(N_{i j+1}^{g_{k}}=1\right)}{\operatorname{Prob}\left(D_{i j}^{g_{k}}=1\right)}\right) \\
& =\sum_{k=1}^{G}\left(\frac{\operatorname{Prob}\left(D_{i j}^{g_{k}}=1\right)}{\operatorname{Prob}\left(D_{i j}=1\right)}\right) R_{j}^{a, a^{\prime} ; t, t^{\prime} ; g_{k}} \\
& =\sum_{k=1}^{G} \operatorname{Prob}\left(D_{i j}^{g_{k}}=1 \mid D_{i j}=1\right) R_{j}^{a, a^{\prime} ; t, t^{\prime} ; g_{k}}
\end{aligned}
$$

Using Proposition 11, one can express the overall retention rate for workers aged 20 to $54, R_{j}^{20,54 ; 1, \infty}$, as a linear combination of more narrowly defined rates. It can, for example,

[^40]be expressed as a linear combination of retention rates that condition on age
\[

$$
\begin{equation*}
R_{j}^{20,54 ; 1, \infty}=\sum_{k=0 \text { to } 30}^{\text {by } 5} \alpha_{j}^{20+k, 24+k} R_{j}^{20+k, 24+k ; 1, \infty} \tag{4.2}
\end{equation*}
$$

\]

where $\alpha_{j}^{20+k, 24+k}$ is the proportion of workers aged 20 to 54 in year $j$ that are $20+k$ to $24+k$ years of age. Therefore, exploring for changes along specific dimensions (such as age) can provide insight on the sources of the new aggregate pattern.

A decomposition exercise is the next logical step. A counterfactual retention rate for period $j, R C_{j}$, can be created by holding the sub-population proportions in equation (4.1) constant at year 1 levels, i.e.

$$
\begin{equation*}
R C_{j}=\sum_{k=1}^{G} \gamma_{1}^{g_{k}} R_{j}^{a, a^{\prime} ; t, t^{\prime} ; g_{k}} \tag{4.3}
\end{equation*}
$$

For example, the age constant counterfactual

$$
\begin{equation*}
R C_{j}^{20,54 ; 1, \infty}=\sum_{k=0 \text { to } 30}^{\text {by } 5} \alpha_{1}^{20+k, 24+k} R_{j}^{20+k, 24+k ; 1, \infty} \tag{4.4}
\end{equation*}
$$

shows how the overall retention rate would have looked over time if the age structure of the workforce had not changed. By contrasting estimates of $R_{j}^{20,54 ; 1, \infty}$ and $R C_{j}^{20,54 ; 1, \infty}$, one can gauge the importance of structural explanations for the changing aggregate job stability patterns. To quantify the importance of such compositional changes, I use a OaxacaBlinder decomposition approach. The difference in retention rates across time, $R_{2}-R_{1}$, can be expressed as

$$
\begin{equation*}
R_{2}^{20,54 ; 1, \infty}-R_{1}^{20,54 ; 1, \infty}=\left[R_{2}^{20,54 ; 1, \infty}-R C_{2}^{20,54 ; 1, \infty}\right]+\left[R C_{2}^{20,54 ; 1, \infty}-R_{1}^{20,54 ; 1, \infty}\right] \tag{4.5}
\end{equation*}
$$

By substituting (4.1) and (4.3) into (4.5), one rewrite this differential in a more commonly known form

$$
\begin{align*}
R_{2}^{20,54 ; 1, \infty} & -R_{1}^{20,54 ; 1, \infty} \\
& =\left[\begin{array}{c}
\alpha_{2}^{g_{1}}-\alpha_{1}^{g_{1}} \\
\alpha_{2}^{g_{2}}-\alpha_{1}^{g_{2}} \\
\vdots \\
\alpha_{2}^{g_{G}}-\alpha_{1}^{g_{G}}
\end{array}\right]^{\prime}\left[\begin{array}{c}
R^{20,54 ; 1, \infty_{2}, g_{1}} \\
R^{20,54 ; 1, \infty_{2}, g_{2}} \\
\vdots \\
R_{2}^{20,54 ; 1, \infty, g_{G}}
\end{array}\right]+\left[\begin{array}{c}
\alpha_{1}^{g_{1}} \\
\alpha_{1}^{g_{2}} \\
\vdots \\
\alpha_{1}^{g_{G}}
\end{array}\right]^{\prime}\left[\begin{array}{c}
R_{2}^{20,54 ; 1, \infty ; g_{1}}-R_{1}^{20,54 ; 1, \infty ; g_{1}} \\
R_{2}^{20,54 ; 1, \infty ; g_{2}}-R_{1}^{20,54 ; 1, \infty ; g_{2}} \\
\vdots \\
R_{2}^{20,54 ; 1, \infty ; g_{G}}-R_{1}^{20,54 ; 1, \infty ; g_{G}}
\end{array}\right] \tag{4.6}
\end{align*}
$$

The first term in equation (4.5) represents the difference that can be attributed to group composition changes. The second term reflects what remains to be explained, i.e. due to changes in sub-group retention rates. ${ }^{3}$

### 4.2.2 Sub-Population Patterns

Proposition 11 can also be used to explore the source of sub-populations retention rates; a sub-population retention rate can be expressed as a linear combination of more narrowly defined rates. But the decomposition exercise, as presented in Section 4.2.1, will not address the counterfactuals of interest. More specifically, it cannot answer the following question, "does compositional change in the workforce play a role in the new (sub-population) retention rate patterns".

An example best illustrates this point. The one year retention rate for workers with less than one year of initial tenure can expressed as a linear combination of rates that condition on the worker having less than one year of tenure and being in a 5 -year age bracket, i.e.

$$
\begin{equation*}
R_{j}^{20,54 ; 1,11}=\sum_{k=0 \text { to } 30}^{\text {by } 5} \alpha_{j}^{20+k, 24+k} R_{j}^{20+k, 24+k ; 1,11} \tag{4.7}
\end{equation*}
$$

[^41]where $\alpha_{j}^{20+k, 24+k}$ represents the proportion of prime aged workers with less than one year of initial tenure in year $j$ that are $20+k$ to $24+k$ years of age. The counterfactual, $R C_{j}^{20,54 ; 1,11}$,
\[

$$
\begin{equation*}
R C_{j}^{20,54 ; 1,11}=\sum_{k=0 \text { to } 30}^{\text {by } 5} \alpha_{1}^{20+k, 24+k} R_{j}^{20+k, 24+k ; 1,11} \tag{4.8}
\end{equation*}
$$

\]

controls for the age structure of workers with less than one year of tenure, and not of the workforce, and as such, cannot separate ageing from tenure effects.

In the following proposition, I propose a refinement to existing decomposition techniques that remedies this limitation.

Proposition 12 Given $G$ age groups, i.e. $\left[a_{1}, a_{1}^{\prime}\right], \ldots,\left[a_{G}, a_{G}^{\prime}\right]$ that are mutually exclusive and exhaust the conditioned population, one can write the retention rate, $R_{j}$, as

$$
\sum_{k=1}^{G}\left(\frac{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1 \mid y_{i j}^{a_{k}, a_{k}^{\prime} ;, t, \bar{t}}=1\right) \operatorname{Prob}\left(y_{i j}^{a_{k}, a_{k}^{\prime} ; t, \bar{t}}=1 \mid y_{i j}^{a, a^{\prime} ; t, \bar{t}}=1\right)}{\sum_{l=1}^{G} \operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1 \mid y_{i j}^{a_{l}, a_{l}^{\prime} ; t, \bar{t}}=1\right) \operatorname{Prob}\left(y_{i j}^{a_{l}, a_{l}^{\prime} ; \dot{t}, \bar{t}}=1 \mid y_{i j}^{a, a^{\prime} ;, t, \bar{t}}=1\right)}\right) R_{j}^{a_{k}, a_{k}^{\prime} ; t, t^{\prime}}
$$

where $\underline{t} \leq t$ and $\bar{t} \geq t^{\prime}$
proof: Based on Proposition 1, one can write

$$
\begin{equation*}
R_{j}=\sum_{k=1}^{G}\left(\frac{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1, y_{i j}^{a_{k}, a_{k}^{\prime} ;, t, \bar{t}}=1\right)}{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1\right)}\right) R_{j}^{a_{k}, a_{k}^{\prime} ; ;, t^{\prime}} \tag{4.9}
\end{equation*}
$$

From Bayes' Law

$$
\begin{equation*}
\frac{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1, y_{i j}^{a_{k}, a_{k}^{\prime} ; t, t, \bar{t}}=1\right)}{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1\right)}=\frac{\operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t^{\prime}}=1 \mid y_{i j}^{a_{k}, a_{k}^{a_{k} ; t, t}, \bar{t}}=1\right) \operatorname{Prob}\left(y_{i j}^{a_{k}, a_{k}^{\prime} ; t, \bar{t}}=1\right)}{\sum_{l=1}^{G} \operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, t t^{\prime}}=1 \mid y_{i j}^{a_{l}, a_{l}^{\prime} ; t, \bar{t}}=1\right) \operatorname{Prob}\left(y_{i j}^{a_{l}, a_{i}^{\prime} ;, t, \bar{t}}=1\right)} \tag{4.10}
\end{equation*}
$$

and

$$
\begin{equation*}
\operatorname{Prob}\left(y_{i j}^{a_{k}, a_{k}^{\prime} ; ;, \bar{t}}=1\right)=\operatorname{Prob}\left(y_{i j}^{a_{k}, a_{k}^{\prime} ; t, \bar{t}}=1 \mid y_{i j}^{a, a^{\prime} ; t, \bar{t}}=1\right) \operatorname{Prob}\left(y_{i j}^{a, a^{\prime} ; t, \bar{t}}=1\right) \tag{4.11}
\end{equation*}
$$

Substitute (4.10) and (4.11) into (4.9)
Using Proposition 12, one can cleanly identify the ageing effect. The retention rate for workers aged with less than one year of tenure, for example, can be written as

$$
\begin{align*}
& R_{j}^{20,54 ; i, 11}= \\
& \sum_{k=0 \text { to } 30}^{\text {by } 5} \frac{\operatorname{Prob}\left(y_{i j}^{20,54 ; 1,11}=1 \mid y_{i j}^{20+k, 24+k ; 1, \infty}=1\right) \operatorname{Prob}\left(y_{i j}^{20+k, 24+k ; 1, \infty}=1 \mid y_{i j}^{20,54, i, \infty}=1\right) R_{j}^{20+k, 24+k ; 1,11}}{\sum_{l=0}^{\text {by to } 30} 5}\left[\operatorname { P r o b } ( y _ { i j } ^ { 2 0 , 5 4 ; 1 , 1 1 } = 1 | y _ { i j } ^ { 2 0 + l , 2 4 + l i , i , \infty } = 1 ) \operatorname { P r o b } \left(y_{i j}^{\left.\left.20+l, 2^{24+l i, i, \infty}=1 \mid y_{i j}^{20,54 ; 1, \infty}=1\right)\right]}\right.\right. \tag{4.12}
\end{align*}
$$

By holding the conditional probability of being in each age group constant at year 1 levels in equation (4.12), i.e.

$$
\begin{aligned}
& R C_{j}^{20,54 ; 1,11}=
\end{aligned}
$$

one can see how the retention rate of newer jobs would have looked if the age structure of the working population had not changed.

The arguments of Proposition 12, and that of equation (4.13), can be generalized to control for other workforce characteristics. In addition to the effects of ageing, Section 4.3.2 will also control the effects of gender, education and industry structure on the stability of workers with less than one year of initial tenure.

Finally, one can quantify compositional effects across time for any sub-population retention rate, $R_{j}$, using

$$
\begin{equation*}
R_{2}-R_{1}=\left[R_{2}-R C_{2}\right]+\left[R C_{2}-R_{1}\right] \tag{4.14}
\end{equation*}
$$

and replacing $R C_{2}$ with the appropriate counterfactual. ${ }^{4}$ The first term in equation (4.14) represents the difference that can be attributed to group composition changes in the workforce, and the second, what remains to be explained.

[^42]
### 4.3 Source of the Aggregate Patterns

In this section, I provide a thorough examination of potential causes for the new aggregate job stability patterns previously documented in Chapter 2. Some structural explanations include: decline of unions, changes in industry structure, decline in seasonal jobs; increasing education levels, increased labour force attachment of women and ageing of the population.

The long term decline in union coverage in Canada is well documented. ${ }^{5}$ This decline could account for the new patterns only if union covered jobs were less stable than non-union jobs. This presumption is counter-intuitive; it contradicts a fundamental goal of unions to safeguard jobs, and is not reflected in the data. Figure 2.11 clearly indicated that union covered jobs were significantly more stable than their non-union counterpart.

Industry structure in Canada has undergone many changes over the last 30 years, both within and across industries. Figures 2.6 and 2.7 showed that the changing job stability patterns were not restricted to individual sectors, but economy-wide. Yet, as Figure 2.7 illustrated, there are important level differences across sectors. As a result, a change in the relative importance of sectors, such as the well-documented decline of the primary sector, could have impacted the overall long term job stability patterns in Canada. A decomposition exercise can help judge the merits of this hypothesis. The LFS uses the NAICS industry classification system which divides the economy into twenty broad categories. Although there are further divisions, the decomposition was restricted to the 20 categories to ensure more accurate industry retention rates. Figure 4.1 indicates that holding the industry structure constant at 1977 proportions does not significantly alter the pattern in the data. Therefore, the changing industry structure, both within and across sectors, does not appear to be the source of the new patterns in job stability.

[^43]

The possibility of Employment Insurance (EI) reforms in 1989 and the mid 1990s contributing to the new aggregate job stability patterns requires some attention; particularly its impact on seasonal work. ${ }^{6}$ Researchers (e.g., Green and Sargent (1998); and Shen (2004)) found seasonal workers to be more responsive to changes than non-seasonal workers. Changes to the program rules could have limited the appeal of seasonal work, encouraging workers to seek more stable forms of employment. That is simply not the case. Marshall (1999) identifies a long term decline in seasonal variation in employment starting in 1976, but that decline flattens in the 1990s. An industry decomposition also indicates that a possible movement towards less seasonal industries is not critical to job stability patterns. Furthermore, Figures 2.9 and 2.10 show that the changes in patterns are not restricted to regions historically dependent on seasonal jobs. The job stability patterns of Ontario, for example, the least seasonally dependent region, mirrors those of Canada. Finally, changes in job stability are not restricted to periods of the year where seasonal jobs are historically most prevalent, i.e. from May to October. The same patterns hold true if the analysis is

[^44]restricted to January data, or that of any other month.
Canadian workers have become more educated over time, and there is a positive correlation between education and job stability (see Figure 2.5). Together, these findings indicate that education may play a role in the new aggregate patterns. Figure 4.2 holds the proportion of workers with grade 10 or less of education constant at 1977 levels. ${ }^{7}$ The results indicate that controlling for education can explain part of the increase in aggregate job stability.


Figures 4.3 and 4.4 provide the same educational counterfactual, i.e. holding the proportion of workers with grade 10 or less of education, but for males and females separately. Controlling for education matters for both. The decline, in absolute terms, is similar across

[^45]gender, albeit slightly larger for males.


Table 4.1 quantifies the importance of education in explaining overall job stability patterns by comparing the 1987-1989 and 1998-2000 periods. The decomposition results are expressed in percentage terms. As discussed in Section 4.3, a compositional effect of $100 \%$ would mean that changes in group composition can fully account for the difference in job stability over time. Table 4.1 shows that controlling for education composition can explain $25.7 \%$ of the increase for male, but only $8.8 \%$ for females. These results should not be taken to mean that education does not matter for females. Given the findings of Figures 4.3 and 4.4, one should conclude instead that for females there are other factors at work; factors related to the important increases in labour force attachment of women also matter.

Table 4.1: Decomposition Across Business Cycles: 1987-88 and 1998-99

| Specification | Compositional <br> effect <br> $(\%)$ | Remaining <br> effect <br> $(\%)$ |
| :--- | :---: | :---: |
| Education Decomposition |  |  |
| $\quad$ Overall | 16.0 | 84.0 |
|  |  |  |
| Male | 25.7 | 74.3 |
| Female | 8.8 | 91.2 |

Interestingly, also controlling for the proportion of workers with university degrees barely impacted the counterfactual. ${ }^{8}$ There are three reasons why this is the case. One, the proportion of workers with university degrees has increased over time, but this highly educated group still remains a minority in the workforce. Two, differences in job stability between adjacent education groups are much less pronounced at the upper end of the education distribution. Three, it can easily be shown that also controlling for the proportion of workers with university degrees only imposes one additional cross group restriction-that the fraction of grade 11 or more educated workers with university degrees remains constant over time. The data does not find this restriction too onerous. Changes to the education question introduced in 1990 preclude any further breakdown of the middle group, i.e. those with at least grade 11, but less than a university degree. ${ }^{9}$ I believe that the same three

[^46]reasons mentioned above also apply to the middle group. As such, controlling only for the proportion of low educated workers will be a good approximation of the true educated effect.

Some economisists/demographers, including David Foot as a leading proponent, have argued that the demographic composition of a society has strong economic implications. Foot and Stoffman (1996) argued that the baby-boom cohort, those born between 1947 and 1966, through its sheer size has changed the economy and will continue to do so as it ages. One possible link between the demographic structure and job stability can be seen through the lens of a search model. Within the Burdett (1978) framework, jobs are inspection goods. Baby-boomers who are now in the latter stages of their careers will have a lower probability of receiving a better outside offer, and as a result, job stability will increase. In addition, labour force participation of women has increased dramatically over the last thirty years. More women now than ever are permanently attached to the workforce.

Figure 4.5 takes into account these two changes by holding both the age of the workforce and gender composition constant at 1977 level. ${ }^{10}$ Starting in the mid 1980s, the age-gender constant counterfactual is consistently lower than the overall rate, confirming the prediction of the theory. The differential becomes economically significant as of 1992, averaging 2.68 percentage points over the 1992-2003 interval.

[^47]

Figures 4.6 and 4.7 separate the baby-boom effect from the increase in women's participation by carrying out an age constant decomposition for males and females separately. By focussing on the male rate, one can more cleanly identify the effect of the demographic changes in the workforce. Starting in the mid 1980s, the male age-constant rate is consistently lower than the age-varying one, with the gap averaging 2.12 percentage points over the 1992-2003 period. More importantly, there is still a break in the pattern in the early 1990s. The demographic changes cannot explain why the male retention rate has stabilized in the second half of the sample. For females in Figure 4.7, it is a similar story. The main difference is that for the 1992-2003 period the gap is larger, averaging 3.3 percentage points. The increased labour force participation of women appears to have re-enforced the baby-boom effect.


Table 4.2 quantifies the importance of ageing and gender in explaining overall job stability patterns by comparing the 1987-1989 and 1998-2000 periods. Table 4.2 shows that controlling for age and gender composition can explain close to half of the increase in overall stability. ${ }^{11}$ I separate the baby-boom effect from the increase in women's participation by carrying out a separate age constant decomposition for males and females. By focussing on the male rate, one can more cleanly identify the effect of the demographic changes in the workforce. Table 4.2 shows that ageing has more explanatory power for men ( $54.7 \%$ for men versus $45.3 \%$ for women) indicating that holding the female age structure constant cannot account for the full impact of the increased labour force attachment of women. Finally, a gender differential counterfactual was also estimated; one where females are given the male retention rate. This gender counterfactual can explain $25 \%$ of the increase in aggregate job stability.

| Table 4.2: Decomposition Across Business Cycles: 1987-88 and 1998-99 |  |  |
| :--- | :---: | :---: |
| Specification | Compositional <br> effect <br> $(\%)$ | Remaining <br> effect <br> $(\%)$ |
| A. Age/Gender Decomposition | 46.8 | 53.2 |
| $\quad$ Overall | 54.7 | 45.3 |
| B. Age Decomposition | 40.8 | 59.2 |
| $\quad$ Male |  |  |
| $\quad$ Female |  |  |
| C. Gender Differential Decomposition |  |  |
| $\quad$ overall | 25.0 | 75.0 |

In a final decomposition, I control for both education and age. ${ }^{12}$ In doing so, one can also account for any multiplicative effect, i.e. that the ageing effect may not be constant across education groupings. Figure 4.8 shows that education and ageing can explain a very large part of the changes to male stability.

[^48]

Table 4.3 shows that the combined effect can account for $86.1 \%$ of the increase in male job stability from the late 1980 s to the late 1990 s, with the multiplicative effect only accounting for $5.7 \%$ of the change.

Table 4.3: Decomposition Across Business Cycles: 1987-88 and 1998-99

| Specification | Compositional <br> effect <br> $(\%)$ | Remaining <br> effect <br> $(\%)$ |
| :--- | :---: | :---: |
| Education and Age Decomposition <br> Male | 86.1 | 13.9 |

### 4.4 Sources of the Low Tenured Worker Pattern

In this section I explore the source of the dramatic increase in stability experienced by workers with less than one year of initial tenure. Given the significance of the change, it is important to search for its exact source. The section is divided into two parts. In the first part, I explore potential structural explanations. I conclude that of the three main factors
that drove the overall results, i.e. gender, ageing and education, only education matters for newer jobs. I also find that much of the change remains unexplained. In the second part, I present a model for better understand these unexplained changes along tenure lines.

### 4.4.1 Compositional Changes

Gender effects do not explain the striking increase in stability of newer jobs. Figure 2.16 shows that this increase in stability at low levels of tenure was present for both males and females, with the relative gap remaining fairly constant over the 1990s and early 2000 s. As indicated in Figure 2.17, the decline in the male-female retention rate differential was observed over longer tenured jobs.

Using the proposed counterfactual method, i.e. equation (4.13), I control for the age structure of the workforce in the retention rate for workers with less than one year of initial tenure. Using similar logic, I also construct age constant retention rates for workers with one or more years of tenure. ${ }^{13}$ Figures 4.9 and 4.10 show the age constant counterfactuals for the two tenure categories. In both cases, the analysis was restricted to males in order to more cleanly identify the effect of a demographic change. As Figure 4.9 indicates, ageing of the workforce has very little impact on newer jobs. For longer term jobs (see Figure 4.10), the impact is slightly stronger with an average drop of 1.1 percentage points over the 1992-2003 interval. Additional tenure breakdowns (i.e. one to three years, three to seven years, seven to twelve years, and twelve and up) reveal that this slight drop was limited to the one to three year group. Therefore, the ageing effect must have had a significant impact on the proportion of workers within each tenure group - a conclusion confirmed in the data.

[^49]
$($


Education, on the other hand, does make a difference. Figure 4.11 shows that holding the proportion of male workers with grade 10 or less constant at 1977 levels, decreases the retention rate in subsequent years. ${ }^{14}$ Interestingly, this dampening effect starts to be significant in the mid 1980s-meaning that the dramatic increase in stability experienced in the 1990s is still very present.


Finally, I also explore the possibilities that technological change, changing industry structure or the rise of non-standard work could have played a role for low tenured jobs. Since the changes in stability of newer jobs are economy wide, the potential source of change must also be wide in scope. Technological change is one such candidate. It has been offered as a determining factor for many changes and could also be at the root of the new patterns in job stability. However, the timing is wrong. The effects of technology would have occurred gradually, beginning well before the 1990s. By contrast, the rise in stability only started in the early 1990s, and the increase was dramatic. Although not relevant to the aggregate level,

[^50]changing industry structure could have mattered for low tenured jobs. Newer jobs account for less than $20 \%$ of the overall workforce, and as such, the impact of such a group could be lost in the aggregate. This is not the case. I controlled for industry decomposition and found its effect negligible. ${ }^{15}$ Changes in the employer-employee relationship is another possible source of change. Vosko, Zukewich, and Cranforid (2003) document that many Canadian are now engaged in non-standard/contingent work. ${ }^{16}$ Non-standard work accounted for $37 \%$ of all jobs in 2002, a rise of $4 \%$ point since 1989. ${ }^{17}$ The new economy literature argues that these types of jobs offer less security, implying that job stability has declined. Yet, this is not the case-job stability actually increased over the 1990s. The timing also precludes other potential links between the two. Since 1993, there has been little, if any, change in the proportion of Canadian workers in non-standard work arrangements. Yet during this same time period, job stability for newer jobs actually began its dramatic climb. ${ }^{18}$ These findings indicate that technological change, changing industry structure and the rise of non-standard work do not appear to drive the results.

### 4.4.2 Employment Insurance Reform

Having examined more obvious options, there remains an intriguing possibility that EI reforms of the early 1990s could be driving the results for newer workers.

The Canadian Employment Insurance (EI) system underwent important eligibility reforms in the early 1990s. Prior to 1990, a worker could be disentitled from receiving EI benefits for a maximum of 6 weeks if he quit without just cause or was dismissed for cause. ${ }^{19}$

Bill C-21 increased that disentitlement period to between 7 and 12 weeks in 1990. In 1993,
Bill C-113 completely eliminated benefits for these two groups. ${ }^{20}$

[^51]There is evidence that these reforms have had an important impact on the labour market. Kuhn and Sweetman (1998) found that the 1993 reform significantly decreased the quit rates of young workers and females. For prime aged males on the other hand, the results were insignificant. Morissette (2004) examined the quit and layoff patterns over the 1983-1999 period and found an economy-wide decrease in quit rates in the 1990s, relative to the 1980s, with no associated change in layoff rates. He also found a relatively larger drop for younger workers. Neither study conditioned on initial tenure which may have masked the localized impact of the reform on prime aged males. Finally, researchers (e.g., Sargent (1998)) have observed an important decrease in the proportion of the unemployed that receive regular EI benefits, and have concluded that EI reforms have played a role in this decline. ${ }^{21}$

There is a precedent in labour economics, within the context of a search framework, to examine the impact of EI reforms on labour market transitions. ${ }^{22}$ But the unemploymentemployment transition is typically the focus, and the benefit rate the parameter of interest. The impact of EI eligibility changes on job stability, however, has remained largely unexplored.

## Match Quality Search Model

The search framework used to examine changes in job stability along tenure lines is in the tradition of the Mortensen-Pissarides class of search models. It is based on the extension by Pries and Rogerson (2005), where imperfect information on jobs is reflected in two ways: one, meeting a potential employer is costly; and two, quality of the employer-employee relationship is match specific and only partially observable prior to match formation. In this world, a job is both an inspection and experience good. After meeting a potential employer, the job seeker observes a signal, $\pi$, about potential match quality (inspection good), with $\pi$ iid drawn from a known distribution $H(\pi)$. Match quality is either good or bad, with $\pi$ reflecting the probability of a good match. Based on this signal, the job seeker must decide whether to form a work relationship with the potential employer or keep

[^52]looking. Only after match formation occurs will more information on quality be revealed (experience good). Revelation takes on an all or nothing process, with quality revealed at rate $\alpha$.

The advantage of introducing an inspection good component to the basic match quality framework (i.e. where a good is only an experience good) is that this allows for government labour market policies to impact job search behavior. A change in policy will affect the quality of matches that are formed and thus job stability.

An employment relationship may end due to a bad match being revealed, or an exogenous job destruction process (occurring at rate $q$ ). The worker displaced by a bad match enters the EI state with probability $\theta$ (and the non-EI state with probability $1-\theta$ ). A job seeker in the EI state receives EI benefits until he finds employment. A job seeker in the non-EI state does not receive any EI benefits for the duration of his unemployment spell, i.e. he has not met the EI eligibility requirements. Job destruction, on the other hand, accounts for shocks that are exogenous to match quality. These could be firm specific, e.g. a firm restructuring, or economy-wide, e.g. a recession. When a job is destroyed, the worker enters the EI state state with probability 1.

The remaining exogenous variables are as follows: $r$ is the interest rate; $\lambda$ represents the arrival rate of meetings; $w(\pi)$ is the instantaneous wage and it takes the form $w(\pi)=\pi$. If the match is revealed to be good, $\pi=1$ and the wage becomes 1 . If, on the other hand, the match is revealed to be bad, $\pi=0$ and the wage becomes $0 ; Z_{u}$ is the instantaneous income of looking for work when on EI; and $Z_{u n}$ is the instantaneous income of looking for work when not on EI. ${ }^{23}$

The discounted expected utility of an employed person, $V_{e}(\pi)$, a job seeker on EI, $V_{u}$,

[^53]and a job seeker not on UI, $V_{u n}$, satisfy
\[

$$
\begin{align*}
r V_{e}(\pi) & =\pi+q\left[V_{u}-V_{e}(\pi)\right]+\alpha\left[\pi V_{e}(1)+(1-\pi)\left(\theta V_{u}+(1-\theta) V_{u n}\right)-V_{e}(\pi)\right]  \tag{4.15}\\
r V_{u} & =Z_{u}+\lambda \int_{\bar{\pi}_{u}}^{1}\left(V_{e}(\pi)-V_{u}\right) d H(\pi)  \tag{4.16}\\
r V_{u n} & =Z_{u n}+\lambda \int_{\bar{\pi}_{u n}}^{1}\left(V_{e}(\pi)-V_{u n}\right) d H(\pi) \tag{4.17}
\end{align*}
$$
\]

where the threshold reservation signals for the two types of job seekers are determined by

$$
\begin{gathered}
V_{e}\left(\bar{\pi}_{u}\right)=V_{u} \\
V_{e}\left(\bar{\pi}_{u n}\right)=V_{u n}
\end{gathered}
$$

A job seeker on EI will form a match only if the drawn signal exceeds $\bar{\pi}_{u}$. Similarly, a job seeker who is not on EI will only agree to a match if his signal exceeds $\bar{\pi}_{u n}$.

Conditions (4.15), (4.16) and (4.17) provide most of the insight on the workings of the model. The discounted expected flow of income for an employed person, $r V_{e}(\pi)$, consists of three terms. The first term is the wage; the second term is the average income from a possible change in employment status due to a job destruction. The third term is unique to match quality models. It represents the average income from a change in status due to match quality revelation. With probability $\pi$ the match will be revealed good, with an expected gain of $V_{e}(1)-V_{e}(\pi)$. With probability $(1-\pi) \theta$ the match is revealed bad, the worker is relegated to the EI state, and his expected gain is $V_{u}-V_{e}(\pi)$. Finally, with probability $(1-\pi)(1-\theta)$ the match is revealed bad, he is relegated to the non-EI state, and his expected gain is $V_{u n}-V_{e}(\pi)$. Therefore, a higher signal not only leads to a higher instantaneous wage, it also means that the match has a higher probability of being revealed good.

The discounted expected flow of income for a job seeker on EI, $r V_{u}$, consists of the EI benefit, $Z_{u}$, and an option value reflecting the expected benefit of continued search. The discounted expected flow of income for a job seeker not on EI, $r V_{u n}$ is similarly defined with one important difference--he faces a lower instantaneous income from searching. Not
surprisingly, his reservation signal will be lower. There are two effects at work: one, he cannot afford to look as long (i.e. $Z_{u n}<Z_{u}$ ), and two, he also accounts for the $\theta$ probability that a future job separation will lead to the EI state.

A closed form solution is unduly complex and not intuitive, and instead I provide a discussion of how a change in exogenous parameters will affect the threshold signals-the main outcomes of interest. The implications that follow are derived from a numerical simulation and were found to hold over a wide range of parameter values. ${ }^{24}$

The model predicts that an increase in the arrival rate of offers, i.e. $\lambda$, will increase both threshold signals. The intuition is straightforward; given that future offers are now expected to arrive at a faster rate, it is now less costly to reject an offer. As a result, the job seeker, whether on EI or not, can now afford to be more selective in his acceptance of offers. ${ }^{25}$

A decrease in future eligibility, i.e. $\theta$, will increase the reservation signals for both types of job seekers. A lower $\theta$ raises the probability that a future job separation (resulting from a match being revealed bad), will end in the non-EI state. The option value of accepting employment is now lower, and a job seeker will require better assurances of match quality before agreeing to a match; this is reflected in a higher reservation signal. ${ }^{26}$ Thus, for a given unemployment state, the job seeker's next job will be more stable. More importantly, an increase in both reservation signals implies that the minimum observed job quality will have increased. It is no longer optimal for the job seeker to form matches will low likelihood of success. One should therefore see fewer jobs ending at very low levels of tenure.

Finally, one would have greater confidence in the appropriateness of a match quality framework if the changes observed in the data were more prevalent at low levels of initial

[^54]tenure. It would be less plausible to believe that match quality information is only revealed over a period of years.

## Results

The rise in stability of newer jobs, both in timing and scope, is consistent with the EI eligibility reforms of the early 1990s. This is supported through the dramatic rise in stability which starts in the period surrounding the reforms. The changes are economy-wide, i.e. across gender (see Figure 10), regions (see Figure 13 and 14), and a host of other dimensions as shown in Table 3.3.

Figure 2.15 shows retention rates that condition on very low levels of initial tenure. This figure provides three important insights: first, the increases in stability are most dramatic in the first few months of initial tenure. From 1993 to 2000 , retention rates for workers with initial tenure of two months or less increased by 20 percentage points. Within the framework of my model, this finding is consistent with a decrease in $\theta$, i.e. the probability of future eligibility, in the sense that the minimum signal quality, i.e. $\bar{\pi}_{n u}$, has now increased. Matches that have little chances of success are not being formed anymore. This explains why the increases in stability are more modest in scale once one conditions on having worked more than a few months. Second, retention rates initially decrease with tenure. In all but one year, the one-year retention rate for workers with one to two months of initial tenure was systematically higher than the retention rate for workers with three to five months of initial tenure. This is consistent with the findings of researchers that have estimated monthly retention rates. ${ }^{27}$ Farber (1994) finds that the retention rate bottoms out at approximately 3 months, and then begins a slow ascent. Farber (1999) argues that the initial decrease and subsequent increase is consistent with a Jovanovic (1979) type match quality model. The worker is less willing to leave a job very early due to the option value of the match-there is a possibility that the match may turn out to be good. The simplified information revelation process required for my analysis, i.e. to account for a job

[^55]being both an inspection and experience good; rules out this possibility. What stands out is that, at very low levels of tenure, match quality is an important consideration. Third, the increase in the retention rate during the expansion of the mid-to-late 1990s is consistent with predictions of the model. ${ }^{28}$ Finally, an increase in the arrival rate of meetings, i.e. $\lambda$, will increase the reservation signal of both types of job seekers. One can therefore expect job stability to increase during an expansionary phase of the business cycle.

### 4.5 Conclusion

The main contribution of this Chapter is to provide a detailed accounting of possible sources of the new job stability patterns documented in Chapter 2. Particular attention is given to the rise in stability experienced in the 1990s both at the aggregate level and also for workers with newer jobs. I draw the following three main conclusions:

One, compositional changes in the workforce matter for the aggregate rate. Using standard decomposition techniques, I show that the changing age structure of the workforce can explain approximately half of the increase in job stability in Canada. I also find that both the increase in labour force attachment of women and the rising educational levels have a role to play. For the latter, it is the decreasing importance of low levels of education, and not the rising proportion of workers with university degrees, that matter most.

Two, compositional changes in the workforce do not drive the new patterns for low tenured jobs. Using decomposition tools developed in this thesis, I show that ageing of the workforce is not a key; neither are gender differences. I do find that controlling for low levels of education does impact the rate for newer jobs, but cannot explain the dramatic increase in the 1990s.

Three, through the lens of a search framework, I show that the increased stability of newer jobs is consistent with an improvement in match quality brought upon by a tightening of the EI eligibility requirements in the early 1990s. Particularly, the changes are localized at very low levels of initial tenure, and diminishes considerably once one factors on having

[^56]worked more than a few months.

## Chapter 5

## Concluding Remarks

The objective of this dissertation was to document Canadian job stability patterns and explore their causes.

In Chapter 2, I provide a sound statistical interpretation of the retention rate approach, one that is easily applicable to (repeated) cross sectional data. The main contribution of this Chapter is to show that existing methods for estimating standard errors are biased and provide a consistent alternative. Using the proposed method one can construct correct t-statistics and properly identify breaks in job stability patterns. Finally, I show using CPS data that the downward bias of existing methods may lead to a spurious identification of changing patterns.

In Chapter 3, I provide a thorough description of job stability patterns in Canada over the 1977-2004 period. The empirical approach is based on the retention rate. Retention rates have the advantage of being less sensitive to job inflows than in-progress measures, but require detailed and consistent tenure data on a regular basis. This study uses the Canadian LFS data, the only North American data set to satisfy those stringent data requirements. Using this rich source of tenure data and tools developed in Chapter 2, I find that job stability in Canada has stabilized at historically high levels since the early 1990s. I also find marked increases in stability for women, and particularly for workers with less than one year of initial tenure. This Chapter also compares Canadian and American job stability patterns. It adds to the existing literature by extending the comparison period to include
the late 1990s and early 2000s. Although U.S. data difficulties make it all but impossible to separate cyclical from secular changes, there is some indication of recent change at low levels of tenure in the United States. As part of future work, I plan to further explore this possibility by using the Survey of Income and Program Participation.

In Chapter 4, I explore the sources of the new job stability patterns identified in Chapter 3. I find that compositional changes in the workforce can explain an important part of the new aggregate pattern. The three main factors are ageing, increased labour force attachment of women, and increasing educational levels. Using decomposition tools developed in this Chapter, I also find that these same three compositional effects are not the cause of the new patterns for low tenure workers. This Chapter uses a match quality framework to further explore the changes at low levels of tenure. Within a match quality framework, the much stricter eligibility requirements introduced in the early 1990s for the Employment Insurance (EI) program should lead to an increase in stability of newer jobs. There is now a greater probability that a future job separation will occur with the worker not having met the EI eligibility requirements. As a result, the job seeker will now require greater match quality assurances before agreeing to an employment relationship; this will lead to fewer low quality (and less stable) matches being formed. The empirical findings of this thesis are consistent with these predictions. Job stability has increased in the period following the EI changes, and are localized at the bottom end of the tenure distribution. As part of future work, I plan to further explore the gender differential in job stability, and in particular, why it has disappeared over time.

## Bibliography

Akerlof, G. A., and B. G. M. Main (1981): "Measures of Experience-Weighted Measures of Employment and Unemployment Durations," American Economic Review, 71(5), 1003-1011.

Autor, D. H. (2003): "Outsourcing at Will: The Contribution of Unjust Dismissal Doctrine to the Growth of Employment Outsourcing," Journal of Labor Economics, 21(1), 1-42.

Baker, M. (1992): "Unemployment Duration: Compositional Effects and Cyclical Variability," American Economic Review, 82(1), 313-321.

Beaudry, P., and D. A. Green (2000): "Cohort Patterns in Canadian Earnings: Assessing the Role of Skill Premia in Inequality Trends," Canadian Journal of Economics, 33(4), 907-936.

Bernhardt, A., M. Morris, M. S. Handcock, and M. A. Scott (1999): "Trends in Instability and Wages for Young Adult Men," Journal of Labor Economics.

Burdett, K. (1978): "A Theory of Employee Job Search and Quit Rates," American Economic Review, 68(1), 212-220.

Cahuc, P., and A. Zylberberg (2004): Labor Economics. MIT Press.
Devine, T. J., and N. M. Kiefer (1991): Empirical Labor Economics: The Search Approach. Oxford University Press, New York.

Diebold, F. X., D. Neumark, and D. Polsky (1997): "Job Stability in the United States," Journal of Labor Economics, 15(2), 206-233.

Farber, H. S. (1994): "The Analysis of Interfirm Worker Mobility," Journal of Labor Economics, 12(4), 554-593.
_- (1998): "Has the Rate of Job Loss Increased in the Nineties," Working Paper 5014, Industrial relations Section Princeton University.

[^57]Francois, P., and J. Roberts (2003): "Contracting Productivity Growth," Review of Economic Studies, 70(1), 59-85.

Freeman, R. B. (1976): "Individual Mobility and Union Voice in the Labor Market," American Economic Review, 66(2), 361-368.

Goldberger, A. S. (1991): A Course in Econometrics. Harvard University Press, Cambridge, Massachusetts.

Gottschalk, P., and R. Moffitt (1999): "Changes in Job Instability and Insecurity Using Monthly Survey Data," Journal of Labor Economics, 17(5), S91-S126.

Gower, D. (1993): "The Impact of the 1990 Changes to the Education Questions on the Labour Force Survey," Staff report, Labour and Household Surveys Analysis Division, Statistics Canada.

Green, D. A., and W. C. Riddell (1997): "Job Duration in Canada: Is Long-Term Employment Declining," in Transition and Structural Change in the North American Labour Market, ed. by M. G. Abbott, C. M. Beach, and R. P. Chaykowski, chap. 1, pp. 8-40. IRS Press, Kingston, Ontario.

Green, D. A., and T. C. Sargent (1998): "Unemployment Insurance and Job Durations: Seasonal and Non-Seasonal Jobs," Canadian Journal of Economics, 31(2), 247-278.

Hall, R. E. (1982): "The Importance of Lifetime Jobs in the U.S. Economy," American Economic Review, 72(4), 716-724.

Heisz, A. (1996): "Changes in Job Tenure and Job Stability in Canada," Statistics Canada Research Paper 95.
(2005): "The Evolution of Job Stability in Canada: Trends and Comparisons with U.S. Results," Canadian Journal of Economics, 38(1), 105-127.

Jaeger, D. A., and A. H. Stevens (2000): "Is Job Stability in the United States Falling? Reconciling Trends in the Current Population Survey and the Panel Study of Income Dynamics," in On the Job: Is Long-Term Employment a Thing of the Past?, ed. by D. Neumark, chap. 2, pp. 31-69. Russell Sage Foundation, New York.

Jones, S. R. G., and P. Kuhn (1995): "Mandatory Notice and Unemployment," Journal of Labor Economics, 13(4), 599-622.

Jovanovic, B. (1979): "Job Matching and the Theory of Turnover," Journal of Political Economy, 87(5), 972-990.

Kuhn, P., and A. Sweetman (1998): "Unemployment Insurance and Quits in Canada," Canadian Journal of Economics, 31(3), 549-572.

Lin, Z. (1998): "Employment Insurance in Canada: Recent Trends and Policy Changes," Canadian Tax Journal, 46(1), 58-76.

Marcotte, D. E. (1995): "Declining Job Stability: What We Know and What it Means," Journal of Policy Analysis and Management, 14(4), 590-598.

Marshall, K. (1999): "Seasonality in Employment," Perspectives on Labour and Income, 11(1), 16-22.

Morissette, R. (2004): "Have Permanent Layoff Rates Increased in Canada?," Statistics Canada, Analytical Studies Branch Research Paper No. 218.

Neumark, D. (2000): "Changes in Job Stability and Job Security: A Collective Effort to Untangle, Reconcile, and Interpret the Evidence," in On the Job: Is Long-Term Employment a Thing of the Past?, ed. by D. Neumark, chap. 1, pp. 1-27. Russell Sage Foundation, New York.

Neúmark, D., D. Polsky, and D. Hansen (1999): "Has Job Stability Declined Yet? New Evidence for the 1990s," Journal of Labor Economics, 17(4), S29-S64.
(2000): "Has Job Stability Declined Yet? New Evidence for the 1990s," in On the Job: Is Long-Term Employment a Thing of the Past?, ed. by D. Neumark, chap. 3, pp. 70-110. Russell Sage Foundation, New York.

Polsky, D. (1999): "Changing Consequences of Job Separations in the United States," Industrial and Labor Relations Review, 52(4), 565-580.

Pries, M., and R. Rogerson (2005): "Hiring Policies, Labour Market Institutions, and Labor Market Flows," Journal of Political Economy, 113(4), 785-810.

Ramey, G., and J. Watson (1997): "Contractual Fragility, Job Destruction and Business Cycles," Quarterly Journal of Economics, 112(3), 873-911.

Riddell, C., and W. C. Riddell (2004): "Changing Patterns of Unionisation: The North American Experience, 1984-1998," in Unions in the 21st Century, ed. by A. Verma, and T. A. Kochan, chap. 11, pp. 146-164. Palgrave Macmillan, London.

Rogerson, R., R. Shimer, and R. Wright (2004): "Search-Theoretical Models of Labor Market: A Survey," NBER Working Paper No. 10655.

Rose, S. J. (1995): "Declining Job Security and the Professionalization of Opportunity," Research report 95-04, National Commission for Employment Policy, Washington D.C.

Sargent, T. C. (1998): "The BU Ratio: Prospect and Retrospect," Department of Finance, Working Paper 98-09.

Shen, K. (2004): "How are Labour Market Transitions Affected by the 1996 Employment Insurance Reform?," Manuscript, University of British Columbia.

Stinson, M. H. (2003): "Technical Description of SIPP Job Identification Number Editing in the 1990-1993 SIPP Panels," LEHD Technical Paper, U.S. Census Bureau.

Swinnerton, K. A., and H. Wial (1995): "Is Job Stability Declining in the U.S. Economy?," Industrial and Labor Relations Review, 48(2), 293-304.
—— (1996): "Is Job Stability Declining in the U.S. Economy? A Reply to Diebold, Neumark and Polsky," Industrial and Labor Relations Review, 49(2), 352-355.

Ureta, M. (1992): "The Importance of Lifetime Jobs in the US Economy, Revisited," American Economic Review, 82(1), 322-335.

Vosko, L. F., N. Zukewich, and C. Cranford (2003): "Precarious Jobs: A New Typology of Employment," Perspectives, 15(4), 16-26.

Wooldridge, J. M. (2002): Econometric Analysis of Cross Section and Panel Data. MIT Press, Cambridge, Massachusetts.

Appendices

## Appendix A: Education Categories

The following rules were applied to determine the three mutually exclusive educational categories

- High school - Individuals with 13 or less years of schooling. This education category also includes individuals with some post secondary education as defined by the LFS. This refers to individuals who attended a post secondary institution, but did not complete the required program.
- Post secondary certificate or diploma - Individuals who completed a post secondary program below a university degree. For the 1977-1989 sample, only formal post secondary institutions where included. For the post 1990 sample, the category was expanded to include trade certificates which do not require high school graduation.
- University degree - Individuals with a university degree or more.


## Appendix B: Data Related Issues

This appendix discusses the problems of heaping, rounding and non-response, and how they were addressed in this paper.

Heaping occurs when respondents rely on memory recall and answer imprecisely. For the in-progress tenure variable, this manifests itself in spikes in its empirical distribution at five year intervals. Although the sample years were chosen to ensure a consistent tenure question over time, the wording varies across the two surveys. U.S. respondents are asked how long they have been working continuously for their present employer, while the Canadian question focusses on when you started. Researchers (e.g., Ureta (1992)) found the former question to be less precise, leading to more problems of heaping, i.e. rounding of answers to 5 -year intervals. A similar conclusion was drawn in this study. There was a strong presence of heaping in the U.S. data, and to a much lesser extent in the Canadian data. The null hypothesis of no heaping was consistently rejected at the $10 \%$ significance level. For the Canadian data, asking the question when you started could have led to spikes every half decade (i.e. in $1980,1985,1990, \ldots$ ). This would have generated spurious trends in the retention rates. However, there was no evidence of this type of heaping in the Canadian data. Some researchers (e.g., Diebold, Neumark, and Polsky (1997); Neumark, Polsky, and Hansen (2000)) have opted for a de-heaping adjustment, while others (e.g., Swinnerton and Wial (1995)) have not. This study follows the latter approach. By choosing tenure boundaries (upon which the retention rate conditions) that do not include a 5 -year multiple, the heaping problem is reduced. The drawback of this approach is that this rules out the construction of a four year retention rate for workers with one year or less of initial tenure.

Rounding difficulties stem from the periodicity of valid entries. The less frequent the periods, the more rounding will be required of respondents. The Canadian tenure data is recorded in months throughout the 1977-2003 period, but for the United States the unit of measurements vary across time, with years as the most common unit. For the 1983, 1987 and 1991 CPS supplements, tenure is measured in years. Respondents are only prompted for the number of months if tenure is less than one year. For the 1996 and 2000 supplements
the entries are more detailed. The respondents can answer in days, weeks, months or years. Only if the respondents answers one or two years, will he/she be prompted for a monthly answer. To ensure consistency across countries, both the LFS and CPS data were rounded to the nearest integer in Section 3.4. For example, the newly created entry of 4 years of tenure will include all answers within the half-open interval from 3.5 to 4.5 years of tenure. For all other sections of this paper, the LFS tenure data was not rounded.

Finally, non-response is an issue that plagues all empirical work, but is particularly problematic in the retention rate approach. Because the estimator is a function of sums of weights for two different years, a time-consistent solution to non-response is critical. The CPS provides supplementary weights that account for non-response, except for the 1991 supplement. ${ }^{1}$ Two different methods were used for the non-response problem. In method 1 , the base weights of all years were multiplied by a re-weighting factor. For each year of data, the sample was divided into 42 gender/race/age cells, $J_{1, t}, J_{2, t}, \ldots, J_{42, t}$ (i.e. 2 gender categories - male or female; 3 race categories - white, black and other; 7 five-year age intervals). Further let

$$
J_{j, t}^{\prime}=\left\{\text { observations in } J_{j, t} \text { with missing tenure entries }\right\}
$$

For individual i $\epsilon J_{j, t}$, the re-scaled weight, $b \tilde{w}_{i t}$, is

$$
\tilde{b}_{i t}=\gamma_{j, t} b w_{i t}
$$

where

$$
\gamma_{j, t}=\left(\frac{1}{1-\left(\frac{\sum^{i \epsilon J_{j, t}^{\prime}} \frac{\sum_{b w_{i t}}^{i \epsilon J_{j, t}} b w_{i t}}{}}{}\right)}\right)
$$

In method 2, only the 1991 data was re-weighted, with supplementary weights used for all other years. I found that the choice of method had very little impact on the retention rates

[^58]estimates. Method 1 is preferred simply because of its consistent approach across all years of data.

## Appendix C: Asymptotic Properties

Corollary 1 Assuming iid samples for each year; independence across years, no change in $\dot{p o p u l a t i o n, ~ a n d ~} \lim _{n_{1}^{D}, n_{k}^{r}}=1$ for $k=1,2, j, j+1$ and $r=D, N$, then $\sqrt{n}_{1}^{D}\left(\left(\hat{R}_{j}-\hat{R}_{1}\right)-\right.$ $\left(R_{j}-R_{1}\right) \xrightarrow{d} N(0, V)$ where $V$ depends on $j$, an integer greater than or equal to 2 .

Case 1: $j=2$

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right)+\phi_{3}^{2} V\left(D_{i 2}\right)+\phi_{4}^{2} V\left(N_{i 3}\right)+2 \phi_{2} \phi_{3} \mu \operatorname{Cov}\left(D_{i 2}, N_{i 2}\right) \tag{5.1}
\end{equation*}
$$

Case 2: $j \geq 3$

$$
\begin{equation*}
V=\phi_{1}^{2} V\left(D_{i 1}\right)+\phi_{2}^{2} V\left(N_{i 2}\right)+\phi_{3}^{2} V\left(D_{i j}\right)+\phi_{4}^{2} V\left(N_{i j+1}\right) \tag{5.2}
\end{equation*}
$$

with

$$
\phi_{1}=\frac{E\left(N_{i 2}\right)}{\left[E\left(D_{i 1}\right)\right]^{2}}, \quad \phi_{2}=\frac{1}{E\left(D_{i 1}\right)}, \quad \phi_{3}=\frac{E\left(N_{i j+1}\right)}{\left[E\left(D_{i j}\right)\right]^{2}}, \quad \phi_{4}=\frac{1}{E\left(D_{i j}\right)}
$$

and $\mu$ is the probability that a random chosen person in the population aged 20 to 54 will be 21 to 54. $n_{1}^{D}$ represents the number of individuals that are aged 20 to 54 in year $j$, and $n_{j}^{N}$ represents the number of individuals that are aged 21 to 55 in year $j$.
proof: For ease of notation let $\hat{N}_{j}=n_{j}^{-1} \sum_{i=1}^{n_{j}} N_{i j}, N_{j}=E\left(N_{i j}\right)$ and $V_{N_{j}}=V\left(N_{i j}\right)$, and define $\hat{D}_{j}, D_{j}$ and $V_{D_{j}}$ in a similar fashion. Finally, let $C_{2}=\operatorname{Cov}\left(D_{i 2}, N_{i 2}\right)$

Case 1: $j=2$

$$
\begin{align*}
& \sqrt{n}_{1}^{D}\left(\left(\hat{R}_{2}-\hat{R}_{1}\right)-\left(R_{2}-R_{1}\right)\right. \\
& =\sqrt{n}_{1}^{D}\left(\left(\frac{\hat{N}_{3}}{\hat{D}_{2}}-\frac{N_{3}}{D_{2}}\right)-\left(\frac{\hat{N}_{2}}{\hat{D}_{1}}-\frac{N_{2}}{D_{1}}\right)\right)  \tag{5.3}\\
& =\sqrt{n}_{1}^{D}\left(\frac{\left(\hat{N}_{3}-N_{3}\right) D_{2}-\left(\hat{D}_{2}-D_{2}\right) N_{3}}{D_{2} \hat{D}_{2}}-\frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1} \hat{D}_{1}}\right)  \tag{5.4}\\
& =\sqrt{n}_{1}^{D}\left(\frac{\left(\hat{N}_{3}-N_{3}\right) D_{2}-\left(\hat{D}_{2}-D_{2}\right) N_{3}}{D_{2}^{2}}-\frac{\left(\hat{N}_{2}-N_{2}\right) D_{1}-\left(\hat{D}_{1}-D_{1}\right) N_{2}}{D_{1}^{2}}\right)+o_{p}(1)  \tag{5.5}\\
& =\phi_{1} \sqrt{n}_{1}^{D}\left(\hat{D}_{1}-D_{1}\right)-\phi_{2} \sqrt{n}_{2}^{N}\left(\hat{N}_{2}-N_{2}\right)-\phi_{3} \sqrt{n}_{2}^{D}\left(\hat{D}_{2}-D_{2}\right)+\phi_{4} \sqrt{n}_{3}^{N}\left(\hat{N}_{3}-N_{3}\right)+o_{p}(1)  \tag{5.6}\\
& \xrightarrow[\rightarrow]{d} N\left(0, \phi_{1}^{2} V_{D_{1}}+\phi_{2}^{2} V_{N_{2}}+\phi_{3}^{2} V_{D_{2}}+\phi_{4}^{2} V_{N_{3}}+2 \phi_{2} \phi_{3} \mu C_{2}\right) \tag{5.7}
\end{align*}
$$

Case 2: $j \geq 3$. The proof is similar to Case 1, with on exception. Since the four components of the test statistics, i.e. $\hat{N}_{j+1}, \hat{D}_{j}, \hat{N}_{2}$ and $\hat{D}_{1}$ are functions of different yearly samples when $j \geq 3$, the covariance term is zero.

Replacing the population moments in (5.1) and (5.2) with corresponding sample analogs generates a consistent estimator for each asymptotic variance; for survey data, one would use the weighted sample analogs. Finally, one can account for population growth using the tilde estimator. See section 2.8 for more details.


[^0]:    ${ }^{1}$ Job inflows refers to the inflow of individuals into the workforce.

[^1]:    ${ }^{1}$ The data requirements in the job stability literature are more severe for two reasons. One, employment spells tend to be much longer than unemployment spells. As a result, the use of duration models would require very long panels. Two, the job stability literature has focussed on long term changes which requires more years of data.

[^2]:    ${ }^{2}$ For example, an in-progress job spell of two months means that the respondent was presently employed at the time of the survey and that he had been with the same employer. for the last two months. The data does not say when the job will end; it could be the next day or in twenty years. All that is known is that the full duration of the job spell will be more than two months.
    ${ }^{3}$ Akerlof and Main (1981) carried out a very intuitive discussion of the issues regarding in-progress jobs. They even suggested an experience-weighted job spell approach, which could be calculated by simply doubling the actual in-progress job spell.
    ${ }^{4}$ Hall refers to the completed job spell as the "eventual tenure". The approach in his paper is first to estimate retention rates. From these rates he projects the additional time the person will be at his or her job. The final step consists of adding this estimated additional time to the actual reported job tenure.
    ${ }^{5}$ In his sample, the tenure supplements were only available at five year intervals. With such data, Hall would only be able to construct 5 -year retention rates. He could therefore only be able to determine the fraction of jobs that ended within the first five years of starting, and not, for example, the fraction of jobs ending after one, two or even three years.

[^3]:    ${ }^{6}$ The 10 -year contemporaneous retention rate estimator for 40 year old workers with 5 years of job tenure in year $t$ is
    $\frac{\left(\frac{\text { number of indiv. aged } 50 \text { with } 15 \text { years of tenure }}{\text { population aged } 50}\right)}{\left(\frac{\text { number of indiv. aged } 40 \text { with } 5 \text { years of tenure }}{\text { population aged } 40}\right)}$
    ${ }^{7}$ Hall was not ignorant of this identifying assumption. He rightfully stated that contemporaneous and historical retention rates would be equivalent if the tenure distribution within age groups remained stable over time. He even made an adjustment for population changes when calculating the retention rates. However, this important weakness of the contemporaneous approach led subsequent cross sectional researchers to focus exclusively on the historical approach. As such, stating the use of a "retention rate approach" implicitly assumes the use of a historical approach where the conditional probability is estimated using two repeated cross sections. I follow the same notational approach for the remainder of this thesis.

[^4]:    ${ }^{8}$ Job security refers to involuntary job loss, while job stability does not make a distinction between voluntary and involuntary job separation.
    ${ }^{9}$ Farber has written a series of papers that have found similar results, all of which use the DWS and focus on the distribution approach.

[^5]:    ${ }^{10}$ The Swinnerton and Wial (1996) reply only included adjustments to their overall retention rates, and not to the disaggregated results.
    ${ }^{11}$ Swinnerton and Wial (1996) included data from 1979. When comparing the 1979-83 and 1987-1991 periods they found that job stability only declined by 0.2 percentage points. In addition to some important compatibility issues associated with the May 1979 supplement, they are still faced with the same problem, a comparison across different stages of the business cycle.

[^6]:    ${ }^{12} \mathrm{~A}$ contingent job is one that lasts only a limited period of time.
    ${ }^{13}$ Farber (1998) recognized the difficulties he faced. He states, "This adjustment, admittedly crude, results in an average upward adjustment in three job loss rates from the 1984-1992 DWS's of about 11 percent. While this procedure is surely not perfect, it is difficult to think of a better feasible alternative."

[^7]:    ${ }^{14}$ see Baker (1992) for a further discussion.
    ${ }^{15}$ An example will better illustrate this point. Assume a constant population over time, and samples of 100,000 and 300,000 individuals in years $t$ and $t+10$, respectively. By construction, the average weight in year $t+10$ will be one-third the size of those in year $t$. As a result, the weighted count approach will still provide a "reasonable" estimate of the number of individual in the population aged 50 that have 15 years of tenure in period $t+10$. This is not the case for equation (2.1). A larger period $t+10$ sample will result in a larger numerator, and thus lead to an upward bias in the retention rate.
    ${ }^{16}$ For example, one less obvious assumption is that population changes due to immigration lead to breaks in tenure spells. See Section 2.8 for more details.
    ${ }^{17}$ Neumark, Polsky, and Hansen (2000) use the weighted count approach for their retention rate estimates, but go back to to the estimator of equation (2.1) for the construction of their standard errors.

[^8]:    ${ }^{18}$ Technically one could say that the CPS also has some panel aspect to it. The CPS interviews a household for 4 consecutive months, then after an 8 month period, the households is re-interviewed for another 4 months. The problem is twofold. First, the CPS follow dwellings and not individuals. If a person moves after being in the survey for two months, the CPS will simply interview the new tenant. Second, the tenure question is not part of the regular monthly questionnaire, so there is no panel aspect to the tenure question.
    ${ }^{19}$ A summary of the changes to the tenure related questions of the PSID can be found in Polsky (1999).

[^9]:    ${ }^{20}$ To explore for changes in job stability, Heisz (2005) used the cross sectional tools proposed by Neumark, Polsky, and Hansen (2000) - tools shown to be imprecise in Section 2.2.

[^10]:    ${ }^{21}$ It is important to recognize that no employment status restrictions are imposed in year 1. One could, for example, have an individual only work in year 2.
    ${ }^{22}$ More generally, an $m$ year retention rate is the probability a worker remains with the same employer an additional $m$ years.

[^11]:    ${ }^{23}$ The use of different indexes of summation in the denominator and numerator of equation (2.4) is meant to emphasize the fact that the estimator uses two different cross sections.
    ${ }^{24}$ The year $j$ sample is iid drawn from $F\left(x_{j}\right)$.

[^12]:    ${ }^{25}$ See Section 2.10 for more details on the construction of t-tests.

[^13]:    ${ }^{26}$ Assuming a constant population does allows for demographic changes over time. For example, a large group of 20 year olds in the population of year 1, will become a large group of 21 year olds in year 2. A constant population assumes one can follow a population cohort over two consecutive years.
    ${ }^{27}$ A similar argument holds true for the CPS in the United States.
    ${ }^{28}$ On the other hand, any change in population will affect $\tilde{R}_{1}$ through adj$j_{1}$, but leave $\hat{\hat{R}}_{1}$ unchanged.
    ${ }^{29}$ A simple example will illustrate this point. Assume a population of size $n$ in year 1 and $n+1$ in year 2 . The population is constant except for the arrival of a new immigrant in year 2. Order the year 2 population such that the new immigrant is last. Using equation (2.3) the retention rate would be $\frac{\sum_{i=1}^{n+1} N_{i 2} / n+1}{\sum_{i=1}^{n} D_{i 1} / n}$. By assumption, the new immigrant will not affect the sum in the numerator. As a result, the true retention

[^14]:    ${ }^{31}$ The proof involves simple algebraic manipulation.

[^15]:    ${ }^{32}$ For the remainder of this section I assume that both representative cross sections are of size $n . \hat{Q}_{1}$ and $\hat{R}_{1}$ will be numerically equivalent, and as a result, one can more clearly compare the variance estimators.

[^16]:    ${ }^{33}$ Since $\hat{R}_{1} \equiv \frac{\hat{N}_{2}}{\hat{D}_{1}}$, a larger retention rate is associated with a higher $\hat{N}_{2}$ holding $\hat{D}_{1}$ fixed.
    ${ }^{34}$ The t-test is also the preferred method in the job stability literature.

[^17]:    ${ }^{35}$ I assume that the cross sections are of size $n$ to more clearly compare the variance estimators. In the Appendix, I allow for different sample sizes and for non-randomness of the sample.

[^18]:    ${ }^{36}$ This follows from the fact that the probability a dummy variable takes on the value one equals the expectation of the dummy variable.

[^19]:    ${ }^{37}$ Heisz (2005) did not use weights when applying the NPH method. Robustness checks for different years and sub-populations indicate that weights do not significantly affect the NPH results.

[^20]:    ** The estimated difference is significant at the $5 \%$ level

    * The estimated difference is significant at the $10 \%$ level

[^21]:    ${ }^{38}$ This pattern was found to be robust for other time periods and other sub-populations.

[^22]:    ${ }^{39}$ An increase in the U.S. workforce combined with a relatively stable sample size will result in an increase in the mean sample weight over time.

[^23]:    ** The estimated difference is significant at the $5 \%$ level

    * The estimated difference is significant at the $10 \%$ level

[^24]:    ${ }^{1}$ See Neumark (2000) for a summary of the U.S. evidence.
    ${ }^{2}$ With the Panel Study of Income Dynamics (PSID), for example, the results have varied depending on how the researcher dealt with measurement error (see Jaeger and Stevens (2000)).
    ${ }^{3}$ Heisz (2005) is the only other researcher to have used the LFS master files to examine long term trends

[^25]:    in job stability in Canada. He concludes that the changes experienced in the 1990 s were only cyclical in nature. Extending the data to include the first half of the 2000 s -a period of continued economic growth, and using more precise tools, leads me to different conclusions.
    ${ }^{4}$ These files were accessed on site at the British Columbia Interuniversity Research Data Centre (BCIRDC). The BCIRDC is run and sponsored by the University of British Columbia, University of Victoria and Simon Fraser University, in collaboration with Statistics Canada.

[^26]:    ${ }^{5}$ It was brought to my attention by Craig Riddell that LFS interviewers only provided this extra information if prompted. To determine the impact of this potential difficulty, I examined the in-progress tenure distribution of seasonal jobs - spells that should not exceed one year. Two difficulties arose. First, the PERMTEMP variable, which identifies seasonal work, was only introduced to the monthly survey in 1997. Second, the seasonal flag is job specific. It is possible for a worker to have consecutive jobs with the same employer, the last one being seasonal, resulting in a tenure entry of more than one year. Thus, I can only put an upper bound on the problem using recent data. Based on the 1997 sample, only a minority of seasonal workers reported tenure in excess of one year.
    ${ }^{6}$ January 1976 was the official start of a new survey design, the 1971 design. That design was for all intents and purposes a new survey. It introduced elements such as rotation groups, UI regions, and a revised questionnaire. Interestingly, the 1971 design was not introduced in 1971. Statistics Canada started running it in May of 1974, and it officially came into use in January 1976. The 19 month parallel run can be interpreted as a calibration period. In its early stages, the new design only covered approximately 35,000 households (like the older 1961 design), but spurred by the need for more sub-provincial data, Statistics Canada expanded the sample size beginning in February 1976. With no phone interviews, Statistics Canada needed to permanently hire more field workers. As a result, the ramping up was spread over a 12 to 13 month period. The increase was not uniform across time or province. It started in urban areas (and PEI), probably because it was easier to do so. In early 1977, the sample size reached 55,730 households. It should be recognized that the 1976 weights were adjusted to reflect these changes.
    ${ }^{7}$ Otherwise, the sample means in the denominator and the numerator of $\hat{R}_{1}$ would not converge to $E\left(D_{i 1}\right)$ and $E\left(N_{k 2}\right)$, respectively, resulting in a biased retention rate estimator. A similar argument holds for the variance estimator.

[^27]:    ${ }^{8}$ At a minimum, a year 1 retention rate will condition on individuals working in year 1.
    ${ }^{9}$ The retention rates are calculated in a forward manner, i.e. the retention rate for period $s$ estimates the proportion of jobs that continue into period $s+1$. As a result, a sample extending to 2004, generates retention rates to 2003 .
    ${ }^{10}$ The break in pattern is formally tested in Table 3.1.

[^28]:    ${ }^{11}$ The one exception is the $40-44$ age group. Their retention rates are similar to those of the 35-39 age group, but systematically lower than for 45-49 year olds. This may give credence to a mid-life crisis hypothesis.

[^29]:    ${ }^{12}$ The 1989 retention rates were excluded because the changes made the estimates unreliable; the numerator and the denominator did not use the same education measures.
    ${ }^{13}$ Gower (1993) studied the impact of the changes to the education questions. He found time inconsistencies across all educational breakdowns including the university degree category.
    ${ }^{14} \mathrm{~A}$ precise breakdown can be found in the appendix.
    ${ }^{15}$ The presence of a positive correlation between education and job stability is irrespective of education breakdowns. For example, workers with very low levels of education, i.e. grade 10 or less, were the least stable.

[^30]:    ${ }^{16}$ Over the 1977-2003 period, Statistics Canada has used different industry classification systems. Starting in 1999, the LFS switched to the 1997 North American Industry Classification System (NAICS 1997) from the 1980 Standard Industry Classification (SIC80). Although recently introduced, the LFS has provided a historical series for the NAICS`1997.

[^31]:    ${ }^{17}$ In this thesis union status reflects union coverage, i.e. it includes both union members and non-union members covered by a collective agreement.
    ${ }^{18}$ See Freeman (1976) for more details.

[^32]:    ${ }^{19}$ In the master LFS files, job tenure is recorded in months, with a top coding of 999 months.
    ${ }^{20}$ The arguments are typically made in terms of hazard functions, but given the retention rate is simply

[^33]:    ${ }^{21} R_{j, j+1}$ is the retention rate for a randomly chosen worker aged $a$ to $a^{\prime}$ with $t$ to $t^{\prime}$ months of tenure in year $j$ or $j+1$. A further discussion this 'averaged' retention rate can be found in Chapter 2.

[^34]:    ${ }^{22}$ Over the 1977-2004 period, Statistics Canada has used different industry classification systems. Starting in 1999, the LFS switched to the 1997 North American Industry Classification System (NAICS 1997) from the 1980 Standard Industry Classification (SIC80). Although recently introduced, the LFS has provided a historical series for the NAICS 1997.

[^35]:    ${ }^{23}$ The retention rates in this section do not condition out the incorporated self-employed and full-time student. The incorporated self-employed cannot be identified in the 1983 and 1987 CPS, nor can full-time students in 1983.
    ${ }^{24}$ More precisely, the January 1983 Occupational Mobility, Training, and Job Tenure Supplement; the January 1987 Occupational Mobility, and Job Tenure Supplement; the January 1991 Job Training Supplement; the February 1996 Displaced Workers, Job Tenure, and Occupational Mobility Supplement; the February 2000 Displaced Workers, Employee Tenure, and Occupational Mobility Supplement; and the January 2004 Displaced Workers, Employee Tenure. A detailed discussion of other data issues, i.e. heaping, rounding and non-response, can be found in the appendix.

[^36]:    ${ }^{25}$ That the American data is not all for the same month will introduce some measurement error, but is not expected to be of serious concern.
    ${ }^{26}$ Robustness checks confirm that the age restriction is the not the driving force behind the American patterns. Removing the age restriction, i.e. including all workers 16 years and up, has a level effect.

[^37]:    ${ }^{27}$ For the 1983-1995 period, Heisz (2005) had found a much smaller gap between the two countries. The discrepancy can be attributed to rounding. Heisz (2005) chose not to round the Canadian data.
    ${ }^{28}$ U.S. tenure data is only available at irregular intervals and Canadian union coverage information is only available since 1997 in the LFS. As a result, 2000 is the only year where one can estimate a 4 -year retention rate for union covered workers in both Canada and the U.S.
    ${ }^{29}$ See Chapter 4 for a more detailed explanation of this decomposition approach.

[^38]:    ${ }^{30}$ As mentioned in the introduction, the tenure problems in the SIPP are only present in the pre 1990 data.

[^39]:    ${ }^{1}$ Heisz (2005) estimated two sets of counterfactuals: While the first controlled for tenure only, the second controlled for both tenure and age. In each set, he estimated the overall counterfactual, and one for males and females separately. Although descriptive in nature, these counterfactuals cannot identify the source of change. Ageing of the workforce, for example, will affect the proportion of workers in each tenure category; Older workers are more stable, and as such, are less likely to be in low tenure jobs. Holding both age and tenure constant will therefore overestimate the true ageing effect.

[^40]:    ${ }^{2}$ The presentation of the tools in Section 4.2.1 differs from that of the literature in two ways. First, I present the arguments in terms of the population retention rate, and not its estimator-as is commonly done in the literature. Two, I formally prove the results.

[^41]:    ${ }^{3}$ Assume the overall retention rate increased by .15 and that the first and second term in equation (4.5) equalled .1 and 0.05 , respectively. One would conclude that the retention rate increased by 15 percentage point, 10 percentage point of which can be explained by ageing of the workforce. Said differently, ageing of the workforce can explain $66.6 \%$ of the increase in aggregate stability.

[^42]:    ${ }^{4}$ The structure of the counterfactual rules out the more traditional representation, i.e. one analogous to equation (4.6).

[^43]:    ${ }^{5}$ See Riddell and Riddell (2004) for more details.

[^44]:    ${ }^{6}$ The possibility that EI reforms could affect match quality is explored in Section 4.4.2.

[^45]:    ${ }^{7}$ As previously discussed, there was a change in the education question starting in 1990. For the 1977-1989 period, I held constant the proportion of workers with 10 or less years of education. Gower (1993) analyzed the impact of the 1990 changes to the education question and did not find any important discontinuity in the $0-10$ grouping. As a robustness check, I repeated the decomposition exercise using the ELEMHS variable (instead of EDUCLEV). For the post- 1990 period, the variable records the highest grade of elementary or high school ever completed, and as such may be more comparable with the pre- 1990 entries. The results are essentially the same. Finally, the 1989 retention rates were excluded because the numerator and denominator are not based on the same question.

[^46]:    ${ }^{8}$ When comparing the 1987-1989 and 1998-2000 periods, the compositional effect increased by less than a percentage point.
    ${ }^{9}$ See Gower (1993) for more details.

[^47]:    ${ }^{10}$ Age was broken down into 7 intervals of five years each (i.e. 20-24, 25-29, 30-34, 35-39, 40-44, 45-49, and 50-54 years of age).

[^48]:    ${ }^{11}$ Age was broken down into 7 intervals of five years each (i.e. $20-24,25-29,30-34,35-39,40-44,45-49$, and 50-54 years of age).
    ${ }^{12}$ The workforce was divided onto 14 categories: $20-24$ years of age with grade 10 or less, 20-24 years of age with more than grade 10 , etc.

[^49]:    ${ }^{13}$ All subsequent decomposition exercises in this sub-section will also be based on the proposed method. As previously mentioned, the method can easily be generalized to control for other characteristics such as education and industry structure, to name just a few.

[^50]:    ${ }^{14} \mathrm{I}$ also controlled for both education and age but this did not materially affect the result.

[^51]:    ${ }^{15}$ The decomposition was restricted to the same 20 industry categories used in Section 4.2.
    ${ }^{16}$ i.e. not not full-time, permanent jobs. This would include self-employment, part-time work and also temporary work.
    ${ }^{17}$ See Vosko, Zukewich, and Cranford (2003) for more details.
    ${ }^{18}$ Self-selection is one such link. The timing issue precludes the possibility that less stable workers moved into self-employment, thereby increasing the retention rate of the remaining paid workers. As a robustness check I also estimated retention rates which included the self-employed; the results were not materially different.
    ${ }^{19}$ The decision whether or not to assign a penalty was at the discretion of the EI case worker.
    ${ }^{20}$ See Lin (1998) for a detailed description of EI reforms over the years.

[^52]:    ${ }^{21}$ The other main culprit is the severity of the recession of the early 1990s.
    ${ }^{22}$ See Rogerson, Shimer, and Wright (2004) for an overview of the search theoretical literature, and Devine and Kiefer (1991) and ? for empirical surveys.

[^53]:    ${ }^{23}$ Assuming that both $\theta>0$ and $Z_{u n} \geq 0$ is sufficient to ensure that bad matches will be dissolved.

[^54]:    ${ }^{24}$ One representative set of parameters includes: $\alpha=0.6 ; \lambda=0.8 ; \theta=0.9 ; q=0.08 ; r=0.004 ; Z_{u}=0.5$ and $Z_{u n}=0.1$ generated $\bar{\pi}_{u}=0.557$ and $\bar{\pi}_{u n}=0.344$.
    ${ }^{25}$ In the basic search model where a job is purely an inspection good, a higher arrival rate increases the reservation wage of the job seeker.
    ${ }^{26}$ Cahuc and Zylberberg (2004) present a similar concept. They present a basic search framework that examines the impact of El changes on new entrants. Within their framework, all individuals that are presently working or that have worked are eligible for EI, but those that have never worked are not. Within this framework an increase in EI benefits will actually decrease the reservation wage of the new entrants. By accepting a job there is a chance of becoming unemployed and getting these higher EI benefits. Cahuc and Zylberberg (2004) refer to this as an eligibility effect.

[^55]:    ${ }^{27}$ The arguments are typically made in terms of the hazard function, but given that the retention rate is simply one minus the hazard, this does not materially alter the argument.

[^56]:    ${ }^{28}$ This increase in stability for newer jobs is also also present across gender, regions and industry to name just a few.

[^57]:    __ (1999): "Mobility and Stability: The Dynamics of Job Change in Labor Markets," in Handbook of Labor Economics, ed. by O. Ashenfelter, and D. Card, vol. 3B, chap. 37, pp. 2439-2483. Elsevier, New York.

    Foot, D. K., and D. Stoffman (1996): Boom, Bust and Echo: How to Profit from the Coming Demographic Shift. Macfarlane Walter and Ross, Toronto, Ontario.

[^58]:    ${ }^{1}$ In the LFS the missing values are imputed.

