The Source of the New Canadian Job Stability Patterns

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Abstract

This paper explores the causes of recent changes in Canadian job stability. Using the Labour Force Survey master files (1977-2010), I find that the increases in job stability first observed in the 1990s were in fact long lasting. Results indicate that compositional changes and the increased job stability of women within age and education groups play important roles in explaining the aggregate job stability patterns that emerge.

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1 Introduction

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 such as temporary work and self-employment (e.g. Autor (2003); Vosko, Zukewich, and Cranford (2003)). A perception exists that, in this New Economy, workers have become as disposal as any other resource—implying that job stability has declined. However, researchers who directly examined this issue found modest (if any) evidence of a long term decline in job stability (e.g. Bernhardt, Morris, Handcock, and Scott (1999); Gottschalk and Moffitt (1999); Heisz (2005); Neumark, Polsky, and Hansen (1999)). Except for Heisz (2005) who used Canadian data, this U.S. based literature faced significant data limitations that made it difficult to differentiate between cyclical and secular changes.[1]

Using Canadian Labour Force Survey (LFS) data, Heisz (2005) concluded that there was no long term drop in job stability. He did, however, find some evidence of increased stability in the mid- to late 1990s. By updating the Canadian evidence into the late 2000s, this paper clearly shows that the changes first observed by Heisz (2005) represent long-term changes in job stability, rather than temporary as was previously believed. This paper also goes a step further and explores the causes of these striking new Canadian job stability patterns, particularly the role played by compositional change.

There is a related literature that explores the link between compositional changes and the dampening of the business cycle (i.e. the "Great Moderation").[2] Jaimovich and Siu (2009) argue that young workers (those under the age of 30) experience much greater employment

[1]This may explain why the U.S. job stability literature never went beyond a basic characterization of job stability patterns, and a possible reason why the belief still persists that the employee-employer relationship has changed (e.g. Autor (2003)).

[2]A decline in output growth volatility was observed in the U.S. (e.g. Kim and Nelson (1999); McConnell and Perez-Quiros (2000)) and other G7 countries (e.g. Blanchard and Simon (2001); Stock and Watson (2003)).

volatility than their older counterparts. As such, their declining importance in the workforce, due to the ageing of the population, is an important key to the reduced cyclical volatility. My paper investigates whether such compositional changes can also explain the changing job stability patterns.

Identifying changes in job stability is important for many reasons. Given that pensions are typically associated with long term employee-employer relationships, a decline in job stability would affect the efficient design of employer-based pension plans, and that of government sponsored saving plans, e.g. RRSPs in Canada and 401Ks in the United States. Understanding changes in job stability is also a necessary first step towards better labour market policies—even if job security is the object of interest.[3] There is often a fine line between a quit and a layoff, e.g. voluntary buy-out packages.[4] As a result, focusing exclusively on layoffs (or quits) may lead to incomplete, or even incorrect, policy recommendations. Finally, economic theory does not typically make a distinction between quits and layoffs: there is a separation when the surplus is gone.[5] As a result, job stability is a natural focus. In a world of incomplete contracts, for example, 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)).

The job stability literature has relied on the retention rate, the probability that a job with a particular employer will last one more period, to proxy for job stability.[6] In order to examine long term changes in job stability using retention rates, it is essential to have detailed and consistent tenure data that is available on a regular basis.

In this paper, I use the Canadian LFS files (1977-2010) as repeated cross sections. The

[3] Job security refers to involuntary job loss, while job stability does not make a distinction between voluntary and involuntary job separation.

[4] More specifically, a firm may offer voluntary buy-out packages as a first step in reducing its workforce. More vulnerable workers may decide to accept this option only because their future with the company looks bleak.

[5] There are some exceptions. Within an efficient turnover model where only non-surplus producing matches are dissolved, McLaughlin (1991) labels a separation as a quit (layoff) if the worker (firm) initiates the unsuccessful wage change.

[6] The advantages of the retention rate approach over in-progress measures are discussed in the next section.

LFS is the only North American data set that satisfies the stringent data requirements of the retention rate approach.[7] A consistent job tenure question has been part of the regular monthly LFS questionnaire since its inception. As a result, a full set of one-year retention rates can be constructed dating back to 1977. Using this full set of retention rates, one can identify new patterns and explore their causes. Due to the many similarities between the Canadian and U.S. labour markets, the LFS may shed light on the changes in job stability not only for Canada, but also for the United States.

This work contributes to the literature in the following ways: First, I examine whether the Canadian job stability patterns first identified by Heisz (2005), and by Green and Riddell (1997), extend into the 2000s—a period that started with a strong economic expansion, but ended in a recession. This part of my paper most closely follows that of Heisz (2005) in that we both rely on the LFS master files and both use a retention rate approach.[8] By extending the period of analysis up to 2010, I can show that the recent changes in job stability pattern, previously believed to be temporary in nature, are longer lasting. I find that the retention rate remained at historically high levels into the late 2000s despite the strong showing of the Canadian labour market. Prior to the mid-1990s, the overall retention rate was clearly counter-cyclical; it almost perfectly mirrored the unemployment rate cycle. The changing

[7] Commonly used panel data sets (in North America) fail to meet these stringent requirements in various ways. One cannot separate out the cohort effect from the period effect in both U.S. national longitudinal surveys. The Panel Study of Income Dynamics (PSID) and the Survey of Income and Program Maintenance (SIPP), suffer from serious measurement error with respect to the job tenure data. Finally, the two main Canadian panels, the Survey of Income Dynamics (SLID) and the Labour Market Activities Survey (LMAS) have limited historical coverage. The SLID only dates back to 1993, while the LMAS has only two short panels, a 1986-1987 panel, and a 1988-1990 panel. The Current Population Survey (CPS) which has been used as repeated cross sections by American researchers (e.g. Neumark, Polsky, and Hansen (1999)) also faces data difficulties. Specifically, the tenure question is not part of the monthly CPS questionnaire, but is only included in select supplements. For example, over the 1980s and 1990s it was only asked in the following supplements: January 1983, January 1987, January 1991, February 1996 and February 1998. See Brochu (2006) for a comprehensive discussion of the strengths and limitations of North American datasets—both panel and repeated cross sections.

[8] Heisz (2005) examined job stability patterns over the 1976 to 2001 period. I extend the analysis up to 2010. Green and Riddell (1997) also used the LFS to examine changes in job stability, but relied on the public access files for 1979-1989, and 1991. In the public access files, tenure was only available by broad categories, making it impossible to directly calculate retention rates. As a result, they had to examine the distribution of in-progress jobs, a method more sensitive to job inflows. More importantly, their data ended just prior to a period of important changes in job stability.

pattern which started in the mid-1990s became more apparent in the 2000s. I also find that the rise in job stability of low tenured workers which was first documented by Heisz (2005) continued into the 2000s. By 2007, the retention rate of workers with less than one year of tenure on a job had exceeded 57%—a historical high, and it remained at high levels even during the 2009-2010 recessionary period.

Second, this paper is the first to systematically explore the sources of change in aggregate job stability.[9] I find that the ageing of the workforce, the increased educational attainment of workers, and the increased job stability of women within age and education groups play important roles in the new aggregate job stability patterns. Ageing and rising educational levels appear to have dampened the amplitude of the retention rate cycle (particularly for males) which is line with the findings of the Great Moderation literature (Jaimovich and Siu 2009). With regard to gender, the evidence points towards a cohort-based explanation. More recent cohorts of women behave more like men along the job retention dimension. This is in line with Baker and Drolet (2010) whose findings suggest that cohort effects may play a role in explaining the declining male-female wage ratio. Surprisingly, I also find that the changing industry structure has little explanatory power when it comes to explaining the new aggregate job stability patterns.

Finally, this paper is also the first to examine the source of change in job stability of workers with less than one year of tenure. The traditional decomposition approach will not address the counterfactual of interest when applied to more narrowly defined rates—like those that condition on low tenure. The decomposition controls for the composition of a *conditional* population (i.e. a sub-group of the workforce), and not of the *unconditional* population (i.e. the overall workforce). As such, it cannot address the source of the change

[9]Although the main focus of Heisz (2005) was to characterize the long term job stability patterns for Canada and to compare them with U.S. findings, he also carried out composition-constant counterfactuals. In all cases, he controlled for initial tenure composition. Yet, tenure is clearly an outcome variable; a change in educational attainment, for example, will probably change the tenure composition. As such, controlling for tenure composition can help better characterize the Canadian patterns, but it cannot answer the key question of this paper: What is the source of change? I elaborate on this issue in Section 4.

in job stability.[10] I propose an adjustment to the traditional decomposition exercises that can separate out compositional effects when conditioning on low tenure. Using the proposed adjustment, I find that education is the only compositional change that matters, but even then much remains unexplained.

The structure of this paper is as follows: Section 2 presents the data set and the empirical strategy which is based on the retention rate approach. Section 3 updates the Canadian job stability evidence into the 2000s. In Section 4, I explore the role of compositional changes in explaining the new aggregate job stability patterns, and the dramatic increase in job stability of low tenured workers. Section 5 further examines the relative increase in female job stability by exploring the role played by cohorts effects. Section 6 examines alternatives explanations for the changed patterns. Finally, Section 7 concludes.

2 Data and Empirical Approach

The main data used in this paper is from the LFS master files. The LFS is a large monthly household survey of approximately 54,000 households per month, with a focus on gathering information about the labour market activities of Canadians. A unique strength of the LFS is the quality and consistency of the 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]?"*[11]

My sample contains individuals between the ages of 20 and 55 in the incoming rotation group over the 1977-2010 period.[12] The LFS follows a rotating panel design, where a house-

[10]For more details see Section 4.2.

[11]While the LFS focusses on *when you started*, the American CPS Respondents are asked *how long* they have been working for their present employer. Researchers (e.g. Ureta (1992)) have found the former question to be more precise, leading to less problems of heaping, i.e. rounding of answers to 5-year intervals. A similar conclusion is drawn in this study; heaping does not appear to be an important problem in the LFS.

[12]Although data is available as of January 1976, the first year was a ramping up year for the survey. As such, the 1976 files are much smaller than in subsequent years. The increase in sample size was not uniform across time or province; it started in urban areas (and Prince Edward Island), probably because it was easier to do so. Even if the 1976 weights were adjusted to reflect these changes, the start year plays a critical role in determining the group proportions for the counterfactual exercises. Considering the 33 years of available

hold 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, in the sense that individuals will enter the sample only once.[13]

The upper age limit accounts for the changes in retirement age (e.g. Milligan and Schirle (2008)). By excluding those above the age of 55, one can focus on quits and layoffs, and not have the results tainted by voluntary retirement.[14] The lower age limit accounts for the small fraction of 15-19 year olds who work, with the majority still in school.

This paper uses a retention rate approach to identify the job stability patterns in Canada. The one year retention rate for workers with characteristics c in year j, R_j^c, is simply the probability that such a worker remains with the same employer for an additional year. There are two advantages to adopting a retention rate approach as opposed to in-progress measures (e.g. in-progress mean tenure).[15] One advantage is 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. The second important advantage is that a well-conditioned retention rate will be less sensitive to job inflows than in-progress job measures. Since Canada has experienced a historically large increase in labour force participation of women and a significant demographic change brought about by the ageing of the baby-boomers, job inflows are an important issue for Canada.[16]

data, the benefits of starting the sample in 1977 outweigh the costs of discarding some information.

[13]The LFS does not follow households if they change dwelling. It is therefore reasonable to assume that workers that have less stable jobs also have less stable living arrangements. As such, the proposed strategy of using all months, but focussing only on incoming rotation groups, will minimize this self-selection problem. However, the LFS survey weights are calibrated using all rotation groups; focussing only on the incoming rotation will therefore make the weights less representative of the population. As a robustness check, I re-estimated the retention rates using the alternative strategy favoured by Heisz (2005), i.e. restricting the sample to surveys that are six months apart, say March and September. It does not materially affect my findings.

[14]As a robustness check, I extended the upper age limit to 64 years of age; it does not affect the key findings of this paper.

[15]The LFS and CPS only measure in-progress job spells. An in-progress job spell of two months, for example, means that the respondent has been with the same employer for the last two months, as of the time of the survey. The data does not say when the job will end. All we know is that the full duration of the job will be more than two months. Therefore, an in-progress mean tenure of, say, 5 years would imply that at the time of the survey, workers had an overage of 5 years of tenure with their present employer.

[16]See Hall (1982), Ureta (1992) and, Diebold, Neumark, and Polsky (1997) for further discussion of the merits of the retention rate approach. It should be noted that the retention rate approach has also been

The retention rates are estimated non-parametrically using cross-sectional data. This paper uses a cross-sectional estimator commonly employed in the job stability literature (Heisz (2005), Neumark, Polsky, and Hansen (1999), Swinnerton and Wial (1995)).[17] The retention rate estimator that conditions on the worker having characteristics c and s periods of tenure as of time j, $\hat{R}_j^{s,c}$, can be written as

$$\hat{R}_j^{s,c} = \frac{\tilde{n}_{j+1}^{s+1,c}}{\tilde{n}_j^{s,c}} \tag{1}$$

where $\tilde{n}_j^{s,c}$ is the sum of the base weights of all individuals with characteristics c who have been employed s periods with the same employer as of period j.[18] The cross-sectional estimator requires that the period j and $j+1$ cross sections be drawn from a population cohort; that the two cross sections select from the same pool of individuals, but at different moments in time.

Brochu (2011) shows that this approach will be consistent if the population changes (due to immigration, emigration or deaths) lead to breaks in tenure spells—a relatively mild identifying assumption. This implies that the worker does not stay with the same employer when he migrates to/from Canada. Finally, the standard errors are constructed using the method proposed by Brochu (2011); a method that accounts for the full variability of (repeated) cross-sectional data. Brochu (2011) shows that existing approaches (e.g. Neumark, Polsky, and Hansen (1999)) tend to underestimate the true standard errors. This can lead the researcher to (incorrectly) conclude that job stability has changed.

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

used to explore other continuation/transition rates of interest. For example, it has been used to explore for changes in Canadian job security using both objective (Morissette 2004) and subjective data (Brochu and Zhou 2009).

[17]This estimation approach has also been used outside of the job stability literature to estimate other continuation/transition probabilities (e.g. the probability of staying unemployed (Baker 1992)).

[18]The literature takes advantage of the fact that base weights of representative cross-sections, like the CPS (U.S.) and the LFS (Canada), sum up to their respective populations. As such, $\tilde{n}_j^{s,c}$ is simply an estimate of the number of people in the population with characteristics c who have been employed for s periods at time j.

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 excluded because the process determining their job tenure spell is very different from (non-military) paid employees.

3 Job Stability Patterns

In this section, I update the Canadian job stability evidence by extending the data until 2010. The retention rates are calculated in a forward manner, i.e. the one-year retention rate for year j estimates the proportion of jobs that continue into year $j + 1$. As a result, a sample extending to 2010 generates one-year retention rates up to 2009.

Figure 1 shows the overall one-year retention rate and the annual unemployment rate for 1977-2009.[19] In the early part of the sample, prior to 1993, the overall retention rate was clearly counter-cyclical; it almost perfectly mirrored the unemployment rate cycle. However, the retention rate increased substantially in the 1990s and stayed at historically high levels (i.e. above 80%) during the 1990s and 2000s. The findings for the 2000s are particularly striking given that the Canadian unemployment rate reached a 34-year low in 2007. The narrow width of the 95 percent confidence interval shows that the change in the aggregate pattern was also statistically significant. What is also worth noting is that despite the heightened job security observed in the 2000s, there remains a cyclical component; the overall rate increased over the 2009-2010 recessionary period, as it had over the two previous recessions.[20]

These findings do not match up with the 'New Economy' literature's belief that the employer-employee relationship has changed—that workers are now disposable just like any other resource. If this were the case, one should have seen a *decrease* in job stability—and

[19] I used Canada's official (annual) unemployment rate for individuals 15 years and up. This result is not sensitive to the choice of age group. I got a similar result using the unemployment rate for individuals 20 to 54 years of age.

[20] See Chan, Morissette, and Frenette (2011) for a detailed exposition of how layoff and hiring rates have changed over these three recessions.

not an *increase* as is observed in the data. These aggregate findings do, however, match up better with those of the great moderation literature in the sense that the counter-cyclical pattern is much more muted.[21]

Figure 2 shows differences in retention rates between men and women. Prior to the 1990s, the male retention rate 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.[22] The fact that there was a important gender gap, and that it narrowed over time, is consistent with the findings of the Canadian job stability literature (e.g. Green and Riddell (1997) and Heisz (2005)). What is novel, however, is that the male-female gap became systematically negative in the 2000s—and for most years the negative gap is also statistically significant. The retention rate for men and women did remain at fairly high levels in the 2000s for both men and women which is indicative that the new patterns identified in Figure 1 are not simply gender driven. The muted cyclical pattern that was present in the aggregate rate can be also be observed in the male rate, and to a lesser extent in the female rate.

An important finding of this paper is the continued rise in retention rates of low tenure workers. Figure 3 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 starting in the 1990s it increased dramatically. The rise began in 1992, and continued into the 2000s with the low tenure rate exceeding 57% by 2007—a historical high. The confidence interval in Figure 3 points to a statistically significant change. Interestingly, the retention remained at high levels even during the most recent recession.

Figure 4 shows the retention rate for jobs with one or more years of tenure, while Figure 5 breaks down longer tenure jobs by (initial) tenure category: 1 to 3 years, 3 to 7 years, 7 to 12 years of tenure and 12 years and up. Figure 4 shows that longer tenure jobs have a

[21] I explore the role of compositional change, such as ageing, in the next section.

[22] The decline in retention rates observed over the expansionary periods of the 1970s and 1980s is also more pronounced for women than it is for men. In section 5 I exploit these cyclical differences to better understand the change for women.

clear counter-cyclical pattern. One observes the same muted pattern in the 2000s as was identified in the aggregate rate. Figure 5 shows that longer tenured jobs tend to be more stable. One can also observe a counter-cyclical pattern for jobs that have 1 to 3, and 3 to 7 years of tenure.

At first glance, the above mentioned changes may not appear very large, but even modest changes in the yearly rate will compound, and in turn significantly impact the probability of observing a longer term job. For example, the aggregate retention rate is on average 2.8% higher in the post-1993 period which implies that the probability of a new job lasting at least five years is now 18% higher.[23]

Table 1 documents the job stability patterns along other dimensions by comparing the 1979-1981, 1987-1989, 1998-2000 and 2006-2008 years, all strong expansionary periods.[24] By comparing retention rates at similar stages of the business cycle, one can put in perspective the importance of changes in job stability.[25] The 2006-2008 period is somewhat less comparable in the sense that the labour market was stronger than in the other three periods—as measured by a much lower unemployment rate. Given the clear counter-cyclical nature of the job stability patterns in the pre-1993 period, one can still draw certain conclusions. If it is truly the case that the job stability patterns have not changed, the 2006-2008 rates should be systematically lower than in the other three periods.

Table 1 shows the overall, male, female, and low tenure rates for each of the four expansionary periods. The findings of the last two expansionary periods confirm the presence of new patterns; the more recent rates are systematically higher (both in an economic and statistical sense) than in the late 1970s or late 1980s. The one exception is the male rate

[23]For an employee-employer relationship to last at least five years it must have survived the first, second, third, fourth and fifth year. As such, the probability that a job lasts at least five years is simply the product of five retention rates. I used the retention rate from Figures 3 and 5 to calculate this probability.

[24]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. For example, $R_{1987,1988}$ is the retention rate for a randomly chosen worker in year 1987 or 1988.

[25]This approach takes into account of 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 limited. It also addresses self-selection problems associated with systematic changes in the composition of the workforce.

for 2006/2007. It is essentially the same as in the late 1970s (although still higher than in the late 1980s). This result is in line with Figure 2 which shows a decline in stability for males in the last few years of my sample—a period where the unemployment rate reached historically low levels.

Panel A of Table 2 shows that the new job stability patterns are broadly based; the higher job stability observed in the 1990s continued into the 2000s irrespective of whether one looks at goods versus services, public sector versus private sector, or by region.[26] These changes are not only statistically significant at the 5% confidence level, but also economically significant. The one exception is the goods sector. The rate in the 2000s is only modestly higher than in the late 1970s; the higher rate of the late 1990s did not continue into the 2000s. A breakdown of the goods sector shows that manufacturing is the main culprit, a sector in which job stability decreased by 5.5% points from the late 1990s to the late 2000s.[27]

Panel B of Table 2 shows that the positive correlation between education and job stability identified by Heisz (2005) is still present in the 2000s.[28] The three education attainment categories 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. A precise breakdown can be found in the Appendix. A positive correlation also exists for finer education breakdowns. For example, workers with very low levels of education, i.e. Grade 10 or less, were the least stable. The results for the low education group are robust to the exclusion of post-secondary dropouts.

Finally, Panel B of Table 2 also shows that the positive age profile present in the first three expansionary periods is still present in 2006/2007. The increasing concave (job stability) age profile, combined with the fact that the job stability pattern of younger workers is

[26] This holds true whether one uses the late 1970s or the late 1980s as the base group.

[27] The decline starts in the mid-2000s. There was also a decrease in the primary sector but it represents a much smaller fraction of the goods sector than manufacturing as indicated by the size of the standard errors.

[28] 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 can be problematic. See Section 4 for more details.

more cyclically sensitive (Brochu (2006)) suggests that ageing of the workforce may have an important role to play in the new Canadian job stability patterns. These results are also in line with the relatively higher hours and employment volatility of younger workers (Gomme, Rogerson, Rupert, and Wright (2005); Jaimovich and Siu (2009)).

4 Sources of Change

In this section, I examine the role played by compositional change in explaining the new job stability patterns documented in Section 3. I start by focussing on the aggregate rate, then examine the retention rate of low tenured jobs.

4.1 Aggregate Job Patterns

Due to its straightforward mathematical representation, the aggregate retention rate, R_j, can be expressed as a weighted average of retention rates for any mutually exclusive and exhaustive G sub-groups (sub-populations)

$$R_j = \sum_{g=1}^{G} \gamma_j^g R_j^g \tag{2}$$

where γ_j^g represents the proportion of the working population in year j that are in group g. One can, therefore, explore for the source of change in the aggregate retention rate by creating a counterfactual retention rate which holds the group proportions, i.e. the γ^g's in equation (2), constant at year 1 levels.

Heisz (2005) is the only other researcher to have carried out composition-constant counterfactuals of Canadian retention rates. His counterfactuals controlled for both age and tenure composition.[29] This exercise is useful if one wants to characterize the job stability patterns, but it does not answer the question of interest: What is the source of change of

[29]He carried out age and tenure composition constant counterfactuals of the overall rate, and also of the male and female rates.

the new job stability patterns? The following will best illustrate this point. Assume, for example, that more educated workers have more stable jobs (which is the case) and that workers are now more educated than before (which is also the case). The overall retention rate will change, but so will the tenure composition. By construction, there will be more workers in longer tenured jobs. Tenure is clearly an outcome variable, and as such, one must not control for it if one wants to explore the source of the change in aggregate job stability.

Canadian workers have become more educated over time, and there is a positive correlation between education and job stability (Table 2). Together, these findings suggest that rising education levels may play a role in the new aggregate patterns.

Figure 6 shows education counterfactuals. Holding for the proportions of workers with Grade 10 or less (low education) at 1977 levels clearly matters.[30] Interestingly, controlling for the proportion of workers with university degrees (high education) barely impacted the counterfactual. There are three reasons why this may be the case. First, the proportion of workers with university degrees has increased over time, but this highly educated group still remains a minority in the workforce. Second, differences in job stability between adjacent education groups are much less pronounced at the upper end of the education distribution. Third, it can be easily shown that 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. I believe that the same three 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 education effect. Finally, I carried out

[30] 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 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.

the same educational counterfactual, but for males and females separately. Controlling for education matters for both groups. The decline, in absolute terms, is similar across gender, albeit slightly larger for males.[31]

Using standard decomposition techniques, one can quantify the impact of the compositional change—like a change in education levels. One can express the retention rates differential, $R_2 - R_1$, as

$$R_2 - R_1 = \left[\sum_{g=1}^{G}(\gamma_2^g - \gamma_1^g)R_2^g\right] + \left[\sum_{g=1}^{G}\gamma_1^g(R_2^g - R_1^g)\right] \tag{3}$$

The first square bracket term in equation (3) represents the differences attributable to compositional changes in the workforce. The second reflects what remains to be explained, i.e. changes in sub-group retention rates. Assume the overall retention rate increased by 0.15 and that the first and second terms in equation (3) equalled 0.1 and 0.05, respectively. One would conclude that the retention rate increased by 15 percentage points, 10 percentage points of which can be explained by compositional changes in the workforce. Said differently, changes in workforce composition would explain 66.6% of the increase in aggregate stability.

Table 3 quantifies the importance of rising education levels in explaining overall job stability patterns by comparing the 1987-1989 and 1998-2000 periods, and also the 1987-1989 and 2006-2008 periods. The former comparison will be the main focus since there is a persisting belief that the labour market changed in the early to mid-1990s (i.e. the New Economy literature). By choosing a period just before (late 1980s) and just after (late 1990s), I can further address the validity of this belief. Although 2006-2008 is somewhat less comparable with other expansionary periods (i.e. the unemployment rate is lower), extending the comparison period puts my 2000s findings in better perspective. It also provides a more thorough analysis of up-to-date trends.

The decomposition results are expressed in percentage terms. Table 3 shows that control-

[31] Results are available upon request.

ling for education composition can explain 24.8% of the increase for males over the late 1980s to late 1990s period, but only 8.6% for females. If one extends the analysis to 2006-2008 one finds that education appears to matter more, but the gender difference remains. These results are an early indication that for females there are other factors at work—other than compositional changes. I address this issue further in the next section.

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) maintain 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. Gender composition of the workforce is another factor that merits attention. Over the last thirty years, there has been a dramatic increase in the labour force participation of women. More women now than ever are permanently attached to the workforce. As such, the changing gender composition may also have a role to play in the new job stability patterns.

Figure 7 takes into account these two changes by holding both the age and gender composition of the workforce constant at 1977 level—where age was broken down into seven intervals of five years each.[32] Starting in the mid-1980s, the age-gender constant counterfactual is consistently lower than the overall rate. The differential becomes economically significant as of 1992, averaging 2.9 percentage points over the 1992-2009 interval.

To separate out the workforce ageing effect from those linked to the increased participation of women, I also carried 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 de-

[32] Accounting for gender composition will not, however, account for the fact that women have become more attached to their jobs. I revisit this issue in Section 5.

mographic changes.[33] Starting in the mid-1980s, Figure 8 shows that the male age-constant rate is consistently lower than the age-varying one, with the gap averaging 2.2 percentage points over the 1992-2009 period. It is a similar story for females. The main difference is that for the 1992-2009 period the gap is larger, averaging 3.5 percentage points.

Table 3 quantifies the importance of ageing and gender in explaining overall job stability patterns. Controlling for age and gender composition can explain close to half of the increase in overall stability over the 1987-1989 and 1998-2000 periods. Table 3 also shows that ageing has more explanatory power for men (54.2% for men versus 41.7% for women) indicating that holding the female age structure constant cannot account for the fact that women are now much more attached to their jobs than ever before. If one compares the 1987-1989 and 2006-2008 periods one finds that ageing plays a larger role. For males, it explains 135.2% of the job stability gap. Therefore, if age composition had not changed, one would have seen a lower male rate over the 2006-2008 period—which is in line with the results of Figure 8.[34]

In a final decomposition, I control for both education and age. 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 8 shows that education and ageing can explain a very large part of the changes to male stability. Table 3 shows that the combined effect can account for 84.6% of the increase in male job stability from the late 1980s to the late 1990s.[35] For women, it represents 56.0% of the increase over the same period. Extending the comparison period to 2006-2008 shows ageing and education accounting for 209.5% of the gap; again implying that job stability would be lower if one controlled for these factors.

The evidence provided in this section indicates that an important part of the increase in aggregate job stability can be explained by a dampening of the job stability cycle; a

[33] This approach cannot, however, deal with potential general equilibrium effects where male jobs could be affected by the increase in labour force participation of women.

[34] For men, the fraction explained by ageing (and also by education) is much larger for the two decade comparison mainly because the change in retention rate that one is trying to explain is relatively smaller. The male rate increased by 3.4% point from 1987-88 to 1998-99 (one decade change), but by only 1.4% point from 1987-88 to 2006-07 (two decade change).

[35] Age (alone) can explain 54.2% of the increase, education another 24.8%, and the remaining 5.6% is the multiplicative effect.

moderation caused by compositional change, i.e. ageing and rising educational attainment. Figures 7 and 8 clearly show that the decline in job stability during periods of economic expansion would have been larger if not for compositional changes. They also show job stability rebounding in all recessions (which includes the 2009-2010 recession). When I compare the 1987-1989 and 1998-2000 periods, two periods at very similar stages of the business cycle, I find little difference in stability once compositional change is accounted for. Finally, the fact that compositional changes over-explains the male gap when one extends the comparison period to 2006-2008 is also in line with the Great Moderation argument. If job stability is truly counter-cyclical one should have seen a lower retention rate in 2006-2008 (as compared to the 1987-1989 period) because of the better performing labour market. This is exactly what one would see if ageing and education composition had not changed.

4.2 Low Tenured Job Patterns

In this section, I explore the source of change in job stability for workers with less than one year of tenure. I want to see whether compositional changes can explain why lower tenured jobs are now much more stable than they were in the 1970s and 1980s. By focussing on a rate that conditions on initial tenure, one can also better understand the channels through which compositional changes affect aggregate job stability. If one thinks, for example, that jobs are experience goods (Jovanovic 1979), then older workers who have more labour market experience may have more (better) information on the quality of potential matches, which should result in fewer newer matches being dissolved; one should therefore see a rise in job stability of workers in low tenured jobs.

As with the overall rate, one can express the retention rate for workers with less than 12 months of initial tenure in period j, $R_j^{1,11}$, as a weighted average of G sub-rates.

$$R_j^{1,11} = \sum_{g=1}^{G} \gamma_j^{g|1,11} R_j^{g;1,11} \tag{4}$$

where $\gamma_j^{g|1,11}$ now represents the proportion (weight) of low tenured workers that are in education group g in year j. These proportions are conditional probabilities; as such, one cannot use the standard decomposition approach to quantify compositional changes.

I focus on the education counterfactual to better illustrate this point. Holding the $\gamma^{g|1,11}$'s constant at year 1 levels will control for the level of education in *low tenured jobs*—and not for the whole workforce which is the counterfactual of interest. I propose an alternative representation that remedies this limitation. By Bayes' Law, one can write the (education) group proportion as

$$\gamma_j^{g|1,11} = \frac{Prob_j(\text{“1 to 11 months of tenure”}|\text{“group } g\text{”})Prob_j(\text{“group } g\text{”})}{\sum_{g=1}^{G} Prob_j(\text{“1 to 11 months of tenure”}|\text{“group } g\text{”})Prob_j(\text{“group } g\text{”})} \quad (5)$$

Equation (5) shows that holding the $\gamma^{g|1,11}$'s constant at year 1 levels would confound education and tenure effects; it controls for a non-linear combination of the two. To control for changes in education composition in the workforce one must keep constant the *unconditional* probability of being in a particular education group, i.e. $Prob(\text{“group } g\text{”})$.[36]

Figure 9 shows that holding the proportion of male workers with Grade 10 or less constant at 1977 levels decreases the retention rate in subsequent years. 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.

Gender effects do not explain the striking increase in stability of newer jobs. The 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 2000s. Using the proposed counterfactual approach, I can control for the age structure of the workforce in the retention rate for workers with less than one year of initial tenure. Figure 9 shows the age-constant counterfactuals for

[36] One replaces $\gamma_j^{g|1,11}$ in equation (4) with

$$\gamma_1^{g|1,11} = \frac{Prob_j(\text{“1 to 11 months of tenure”}|\text{“group } g\text{”})Prob_1(\text{“group } g\text{”})}{\sum_{g=1}^{G} Prob_j(\text{“1 to 11 months of tenure”}|\text{“group } g\text{”})Prob_1(\text{“group } g\text{”})}$$

low tenured workers. The analysis was restricted to males in order to more cleanly identify the effect of a demographic change. As Figure 9 indicates, the ageing of the workforce has very little impact on newer jobs.

Table 4 quantifies the importance of rising education levels and ageing in explaining job stability patterns of low tenured workers by comparing the 1987-1989 and 1998-2000 periods, and also the 1987-1988 and 2006-2007 periods. Results are very similar across the two comparison periods. The findings of Table 4 support the claim that ageing has little impact on the stability of low tenured jobs. It explains at most 3.7% of the change for males, and only slightly more for females. Education does matter a bit more. For males, it explains approximately 20% of the rise in job stability. The results of Table 4 also confirm the points made above that much of the increase remains to be explained.

Focussing on equation (4) provides insight as to why ageing has so little effect on the retention rate of low tenured jobs. Controlling for ageing reallocates the weights attached to each conditional retention rate. The impact of this reallocation will depend on the age profile of the low tenured retention rate. A relatively flat age profile will dampen the impact of ageing. This is confirmed in the data. I find that when it comes to newer jobs (jobs that just started) there is very little difference among age groups; older worker are only slightly more likely to stay on with the same employer one more year (as compared to younger workers).[37]

Given that ageing affected the overall rate, it must mean that it must have affected the stability of longer term jobs of the composition within each tenure group.[38] If one controls for the age structure of workers with more than one year of initial tenure the impact is

[37] There is, however, an educational profile for newer jobs which explains why education does impact the retention rate of newer (low tenured) jobs.

[38] This can be more easily seen by assuming only two age groups (a_1 and a_2) and two tenure groups (t_1 and t_2). The overall rate can then be written as $R_j = \gamma_j^{t_1} R_j^{t_1} + \gamma_j^{t_2} R_j^{t_2}$ where

$$\gamma_j^{t_k} = \gamma_j^{t_k|a_1}\gamma_j^{a_1} + \gamma_j^{t_k|a_2}\gamma_j^{a_2}, \quad R_j^{t_k} = \frac{\gamma_j^{t_k|a_1}\gamma_j^{a_1}}{\gamma_j^{t_k|a_1}\gamma_j^{a_1} + \gamma_j^{t_k|a_2}\gamma_1^{a_2}} R_j^{t_k,a_1} + \frac{\gamma_j^{t_k|a_2}\gamma_j^{a_2}}{\gamma_j^{t_k|a_1}\gamma_1^{a_k} + \gamma_j^{t_k|a_2}\gamma_j^{a_2}} R_j^{a_2,t_k}$$

From the above equation one can see that controlling for ageing, i.e. replacing $\gamma_j^{a_k}$ with $\gamma_1^{a_k}$, will affect both the tenure compositions and also the retention rates that condition on different levels of tenure.

slightly stronger. More refined tenure breakdowns (i.e. one to three years, three to seven years, seven to twelve years, and twelve and up) reveal that the drop was limited to the one to three year group. Therefore, the ageing effect must also have impacted the proportion of workers within each tenure group—a conclusion confirmed in the data.

5 Further Exploration of the Female Retention Rate

It well documented that the labour force participation rate of women increased dramatically over the last thirty years, and that the labour market outcomes of women, once they have entered the labour (or work) force, now more closely resemble those of men (e.g. Fortin (2009); Baker and Drolet (2010)). There is also evidence that cohort effects matter for some of these outcomes (e.g., life cycle profile of the female participation rate (Schirle 2008); male-female wage gap (Baker and Drolet 2010)). In this section, I want to better understand the relative increase in the female retention rate by exploring the role played by cohort effects. I want to see whether the changes were restricted to newer cohorts or were experienced by all cohorts of women.[39]

Figure 10 shows the female retention rates by age group.[40] The post-1993 period is clearly more stable for all age groups which indicates some commonality of change across cohorts. However, the dramatic increase in female labour force participation rate combined with the fact that the life-cycle profile of more recent cohorts of women tend to be flatter and have a higher starting point as compared to older cohorts (e.g. Schirle (2008)),[41] point to a possible cohort explanation for the declining male-female retention rate gap.

Figure 11 shows the female retention rate by 5-year birth cohorts. One can see some evidence of a cohort effect. Although a bit noisy, the more recent cohorts appear to have

[39] In section 4, I showed that the changing male-female retention rate differential is not due to compositional changes. For example, ageing and rising educational attainment can account for the majority of the increase in the male retention rate (i.e. 84.6%) from the late 1980s to the late 1990s, but only 56.0% of the (larger) increase in the female rate.

[40] A weighted average of two consecutive years was used to smooth out the series. I used a similar approach in the subsequent cohort approach.

[41] Using the LFS master files, I found that the profiles started to flatten beginning with the 1953-57 cohort.

more stable jobs over their life-cycle.[42] The age profile of the 1938-42 cohort, for example, is systematically lower than for other cohorts. One can also see a systematic gap between the 1948-52 and 1958-62 cohorts, at least in the early part of the age profile (between the ages of 25 and 39). In Figure 12, I present the male evidence. Contrary to women, one cannot see any clear profile differences across male cohorts. Finally, Figures 13 and 14 combine the male and female findings (in the same picture) to better contrast the gender differences, and better show how they have evolved across cohorts. There are systematic gender differences for the earlier cohorts (Figure 13) that are not present for the later ones (Figure 14).[43]

There was a clear cyclical pattern in the female retention rate in the first half of my sample period (Figure 2). Combined with the fact that different cohorts experienced a (particular) recession at different age, it is possible that what one observes in Figure 11 is not just a cohort effect (i.e. it is confounded by the business cycle effect). As a robustness check, I used the fact that female retention rates fell at a faster rate than for males in the late 1970s and late 1980s expansionary periods to identify the cohort effect. More precisely, I compared the male-female differential across cohorts for the 1977-1980, and 1984-1989 periods. The differential is economically significant (in both periods) for earlier cohorts, but declines with subsequent cohorts. By the time one reaches the 1953-57 cohort, the differential is already economically insignificant; the difference (when averaged over each period) is approximately one percentage point. This is further evidence that cohort effects have a role to play in the declining male-female retention rate ratio.[44]

[42] I actually estimated retention rates for 11 birth cohorts (i.e. for the 1928-32, 1933-37, 1938-42, 1943-47, 1948-52, 1953-57, 1958-62, 1963-67, 1968-72, 1973-77, and 1978-82 cohorts). I chose to only present the results of four cohorts (that are ten years apart from each other) to simplify the analyse of the graph. The pattern would be the same if I were to use all 11 cohorts.

[43] As a robustness check, I tried a regression-based approach similar to that of Beaudry and Green (2000). I regressed the retention rate for a given (5-year) cohort group in a given year on a quadratic in age, a quadratic in cohort entry year, an interaction of the linear age and the cohort term, and the unemployment rate. Following Beaudry and Green (2000) I used the quadratically detrended unemployment rate for males 45 to 54 years of age, and instrumented for it using the detrended U.S. unemployment rate. I also tried a specification where I interacted the square of the age with the cohort term to allow for a more flexible adjustment process across cohorts. For the male regression, the cohort terms were all economically and statistically insignificant. For women, cohort effects were present. Both specifications generated age profiles that were generally higher for more recent cohorts.

[44] As an interesting side note, I also explored whether the presence of young children impacted the female

6 Robustness Checks

The long term decline in union coverage in Canada is well documented (e.g. Riddell and Riddell (2004)). 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. Unfortunately, a union question was not part of the regular LFS questionnaire prior to 1997, so one cannot comment on potential long term changes. But over the 1997-2010 period, jobs were more stable in the union sector, with a gap ranging from 12 to 17 percentage points for all jobs, and 12 to 21 for newer jobs. As such, it is very unlikely that the decline in unionization could explain the new job stability patterns in Canada.

The industry structure in Canada is another area that has undergone many changes over the last 30 years, both within and across industries. Table 2 shows that the changing job stability patterns were not restricted to individual sectors, but are economy-wide. However, as Table 2 illustrates, 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 North American Industry Classification System (NAICS) which divides the economy into twenty broad industry categories.[45] Although there are further divisions, the decomposition was restricted to twenty categories to ensure more accurate industry retention rates. Holding the industry structure constant at 1977 proportions, however, does alter the pattern in the data. This holds true for both the aggregate and newer job patterns.[46] Therefore, the

retention rate and whether the link had changed over time due to changing government policies (such as more generous maternity benefits). The presence of small children does not appear to drive any of my findings. In fact, the retention rate of women with children four years or younger has been systematically higher than for other women since 1977.

[45]Although there has been some changes in industry classification systems over time (e.g. from SIC to NAICS), the LFS provides a consistent industry affiliation variable based on the NAICS that goes back to 1976.

[46]The industry-constant counterfactual slightly diverges from the true rate in the last couple of years. This is probably due to the fact that job stability in the manufacturing recently fell. The results are available

changing industry structure, both within and across sectors, does not appear to be the source of the new patterns in job stability.

It is possible that the long term decline in seasonal jobs could explain the rise in job stability—particularly for newer jobs. 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, Table 2 shows 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 as a whole. 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 restricted to January data, or that of any other month.

Changes in the employer-employee relationship is another possible source of change. Vosko, Zukewich, and Cranford (2003) document that many Canadians are now engaged in non-standard/contingent work, i.e. not full-time, permanent jobs. This would include self-employment, part-time work and also temporary work. Non-standard work accounted for 37% of all jobs in 2002, a rise of 4% point since 1989. 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 and remained at high levels into the 2000s. 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. 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

upon request.

the self-employed; the results were not materially different. These findings indicate that the decline in unionization, changing industry structure and the rise of non-standard work do not appear to drive the results.

I now try a different identifying strategy to further investigate the potential role played by ageing and rising educational attainment of the workforce in explaining the new job stability patterns. Specifically, I create two-month LFS panels and carry out regression analysis where the focus is not just the probability of continuing with the same employer, but also the probability of quitting or being laid off.[47] By focussing on the reason for the separation, I can also better understand how ageing and the rising educational of workers affects job stability.

Following Brochu and Green (2011), I take advantage of the rotating panel design of the LFS.[48] Since, individuals remain in the sample for six consecutive months, I can link consecutive months of the LFS to create two-month mini panels covering the 1977-2010 period.[49] I limit my sample to individuals that are employed in period 1 of the mini panel, and that are between the ages of 20 and 54, except full-time students, the self-employed and those in the military, as was done in the main body of my paper. I further restrict my sample to March and September data. By restricting my sample to these months I can ensure that individuals enter the sample only once. I tried different 2 month combinations (that are 6 month apart) and got similar results.

The first column of results in Table 5 shows the marginal effect from a Linear Probability Model on the probability of continuing with the same employer.[50] The dependent variable is

[47] The econometric approach used in the main body of the paper relied on a synthetic cohort approach. As such, it only required the use of repeated cross sections to construct one-year retention rates. This synthetic cohort approach cannot, however, be used to construct one year layoff and quit rates (see Brochu (2011) for more details on the identifying assumptions of the estimator).

[48] Other researchers that have used such an approach include Jones and Riddell (2006), Chan, Morissette, and Frenette (2011), and Campolieti (2011).

[49] The LFS may have panel features, but it is not a panel data set per se; the LFS is officially designed to produce cross sectional samples. As such, it follows dwellings, and not individuals (households). If an individual changes dwelling, he is out of the reach of the survey. It is therefore reasonable to assume that workers that have less stable jobs also have less stable living arrangements. I can minimize this self-selection problem by focusing on one-month transitions.

[50] For Table 5 and the subsequent table, I rely on weighted standard errors where the weights sum to 1 in

a binary variable that equals 1 if the period 1 worker continued with the same employer in period 2 of the mini-panel. The set of (period 1) controls include a female dummy, age, age squared, binary variables for low (grade 10 or less) and high (bachelors and up) education, region dummies, and month and year dummies.

The results strongly support the main findings of my paper. There is a positive and concave relationship between age and job stability, and job stability increases with education. In addition, the marginal effects of age and education are smaller when one focusses on workers with less than one year of tenure (bottom panel of Table 5). Given that there was a larger absolute increase to be explained, these results suggest that age and education have less explanatory power when it comes to low tenured jobs.[51] As a robustness check, I carried out the estimation by gender, and by adding industry controls; neither materially affected my findings.

In the remaining columns of Table 5, I focus on the probability of leaving employment, and by reason. For the second column of results, the dependent variable equals 1 if the worker is no longer working in the next period and answered he quit when asked about why he separated from his last job. The main problem with this definition of quits is that it will not account for those who separate from a job and find a new job before the following month. I, therefore, use a second definition of quits (quit2) which includes those captured under the first definition and anyone who was on a job in the first period and employed but on a different job in the second period.[52] The final binary dependent variable is layoff, which equals 1 if the worker is not working in the second period, and responded that he was laid

each year. It should also be noted that using a probit model or a linear probability model gives very similar results.

[51]The results are not perfectly comparable because I focus on one-month continuations (as opposed to 1-year continuations), and by including time dummies I am assuming a common business cycle effect across groups. I did, however, repeat the same estimation for the strong expansionary periods, the same periods covered in Tables 3 and 4, and got very similar results.

[52]The LFS did ask job-to-job switchers why they left their first job, but only over the 1999-2005 period. 68% of workers who transit directly to a new job by the time of the next monthly survey reported that they quit their previous job. Thus, the second quit definition mainly captures actual quitters. It does, however, also include a significant number of layoffs. It turns out that the results for both measures of quits are qualitatively similar.

off from the previous job.

Both age and education work in the same direction when it comes to quits and layoffs; there is less chance of a transition if the worker is older or more education. What is interesting to note, however, is that their relative importance varies by the reason for the separation. For layoffs, the education profile is much steeper. On the other hand, age appears to matter more for quits. This last result would seem to indicate that the effect of ageing on retention rates is mainly through quits; a result that is in line with the simple Burdett (1978) framework, where jobs are inspection goods.

As a final robustness check, I further examine the link between the retention rates and the Great Moderation (as measured by the unemployment rate). It is possible, for example, that the prolonged boom experienced over much of the 1990s and 2000s changed the nature of the transitions. One could, however, see this going either way - with a better job market people may be more willing to quit, but firms may also be less willing to layoff workers because of fears of refilling the position.

Table 6 shows the correlation between the transitions (as presented in Table 5) and the provincial unemployment—with and without controls.[53] Focussing on the probability of staying with the same employer (retention), one can see that for all workers (all tenure) there is a counter-cyclical pattern, i.e. job stability decreases in periods of expansion, and the counter-cyclical pattern is more pronounced when one adds controls, like age and education. This result strongly supports the findings of Section 4.1, and more precisely Figure 7. Carrying the same analysis, but by initial tenure is also very revealing. It shows that the job stability pattern of low tenured workers is pro-cyclical, but clearly counter-cyclical when one conditions on longer tenure. Therefore, the counter-cyclical pattern of the aggregate rate comes from longer tenured jobs "disappearing" less in bad times.

Focussing on quits and layoffs provides some insight on this finding. The results show that the probability of quitting increases in periods of expansion—when there is a better job

[53]The regression with controls include controls for gender, age, age squared, and dummy variables for education (low and high) and provinces.

market. The probability of a layoff, on the other hand, rises with the unemployment rate, but the strength of this link diminishes with length of service with an employer. As such, these results indicate that job stability increases in periods of recession because the quit effect dominates for longer tenured jobs; they are less likely to quit because of the worsening job market.

7 Conclusion

There are three main contributions to the literature found in this paper: First, I update the job stability evidence into the 2000s. I show that the increases in job stability first observed in the 1990s were long lasting, and not temporary as previously believed. I also show that the rise in job stability of low tenure workers which had started in the 1990s, continued into the 2000s. By 2007, the retention rate for workers with less than one year of tenure had exceeded 57%—a historical high. These findings do not match up with the 'New Economy' literature's belief that workers are now disposable just like any other resource. In fact, this paper shows the opposite: The employer-employee relationship is now more stable than it was in the 1970s and 1980s.

Second, compositional changes and the increased job stability of women within age and education groups matter for the aggregate rate. Using standard decomposition techniques, I show that ageing and rising educational attainment appear to have dampened the amplitude of the retention rate cycle which is in line with the findings of the Great Moderation literature. With respect to education, it is the decreasing importance of low levels of education, and not the rising proportion of workers with university degrees, that matter most. Compositional changes cannot explain why women behave more like men along the retention rate dimension. There is evidence, however, of a cohort effect, as more recent cohorts of women are more attached to their jobs.

Third, compositional changes in the workforce do not drive the new patterns for low tenured jobs. Using decomposition tools developed in this paper, 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.

A Appendix

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

- High school: Individuals with 13 years or less 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 (e.g. degree, diploma).
- 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.

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Table 1: Retention Rates Across Business Cycles

	$R_{1979,1980}$	$R_{1987,1988}$	$R_{1998,1999}$	$R_{2006,2007}$
Overall	.783	.777	.820	.809
	(.002)	(.002)	(.002)	(.002)
Male	.795	.785	.819	.799
	(.002)	(.003)	(.003)	(.003)
Female	.766	.767	.821	.820
	(.003)	(.003)	(.003)	(.003)
Tenure $<$ 1 year	.476	.444	.546	.569
	(.004)	(.004)	(.005)	(.005)

Notes. 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. For example, $R_{1987,1988}$ is the retention rate for a randomly chosen worker in year 1987 or 1988. Standard errors are in parentheses.

Table 2: Retention Rates Across Business Cycles, by industry, region, age and education

	$R_{1979,1980}$	$R_{1987,1988}$	$R_{1998,1999}$	$R_{2006,2007}$
		Panel A		
Sector				
Goods	.778	.784	.828	.790
	(.003)	(.004)	(.004)	(.004)
Services	.786	.774	.818	.816
	(.002)	(.002)	(.002)	(.002)
Agriculture, Forestry, Fishing and Hunting	.662	.639	.640	.615
	(.014)	(.016)	(.016)	(.018)
Mining, Oil and Gas Extraction, Utilities and Construction	.720	.741	.773	.781
	(.007)	(.007)	(.009)	(.008)
Manufacturing	.813	.815	.863	.809
	(.004)	(.005)	(.005)	(.006)
Public	.882	.869	.923	.924
	(.004)	(.004)	(.005)	(.005)
Private	.747	.745	.788	.774
	(.002)	(.002)	(.002)	(.002)
Region				
Atlantic	.760	.753	.769	.797
	(.008)	(.008)	(.008)	(.009)
Quebec	.804	.774	.835	.830
	(.004)	(.004)	(.004)	(.005)
Ontario	.808	.794	.832	.819
	(.003)	(.003)	(.003)	(.003)
Prairies	.719	.750	.798	.775
	(.005)	(.005)	(.005)	(.005)
British Columbia	.761	.776	.817	.794
	(.006)	(.007)	(.006)	(.006)
		Panel B		
Age				
20 to 29	.699	.679	.698	.692
	(.003)	(.003)	(.004)	(.004)
30 to 39	.812	.809	.836	.813
	(.004)	(.004)	(.004)	(.004)
40 to 54	.862	.848	.878	.868
	(.004)	(.004)	(.003)	(.003)
Education				
High school or less	.768	.754	.801	.763
	(.002)	(.002)	(.003)	(.003)
Post sec. diploma/certificate	.820	0.815	.824	.831
	(.006)	(.005)	(.003)	(.003)
University degree	.829	0.823	.856	.845
	(.006)	(.005)	(.005)	(.004)

Notes. 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. For example, $R_{1987,1988}$ is the retention rate for a randomly chosen worker in year 1987 or 1988. Standard errors are in parentheses.

Table 3: Decomposition Across Business Cycles- Aggregate Pattern

	1987-88 and 1998-99		1987-88 and 2006-7	
	Compositional effect (%)	Remaining effect (%)	Compositional effect (%)	Remaining effect (%)
Education Decomposition				
Male	24.8	75.2	63.2	36.8
Female	8.6	91.4	21.9	78.1
Overall	15.7	84.3	32.2	67.8
Age and Gender Decomposition	47.1	52.9	69.9	30.1
Age Decomposition				
Male	54.2	45.8	135.2	-35.2
Female	41.7	58.3	48.2	51.8
Education and Age Decomposition				
Male	84.6	15.4	209.5	-109.5
Female	56.0	44.0	76.2	23.8
	1987-1988	1998-1999	2006-2006	
Prop. of low educated workers (%)	18.7	8.9	5.7	
Prop. of 20-29 year old workers (%)				
Male	33.9	24.8	25.2	
Female	34.8	23.7	23.4	

Table 4: Decomposition Across Business Cycles- Newer Jobs

	1987-88 and 1998-99		1987-88 and 2006-7	
	Compositional effect (%)	Remaining effect (%)	Compositional effect (%)	Remaining effect (%)
Education Decomposition				
Male	20.6	79.4	19.4	80.6
Female	10.9	89.1	6.8	93.2
Overall	16.8	83.2	14.8	85.2
Age and Gender Decomposition	6.1	93.9	5.7	94.3
Age Decomposition				
Male	2.1	97.9	3.7	96.3
Female	10.4	89.6	4.9	95.1
Education and Age Decomposition				
Male	20.9	79.1	20.4	79.6
Female	24.4	75.6	12.9	87.1
	1987-1988	1998-1999	2006-2006	
Prop. of low educated workers (%)	19.6	10.0	6.8	
Prop. of 20-29 year old workers (%)				
Male	55.5	45.1	45.2	
Female	50.6	43.5	43.0	

Table 5: Linear Probability Model: Retention, Quits and Layoff

All tenure				
	Retention	Quit	Quit2	Layoff
Female	-.0008	.0050	.0032	-.0035
	(.0004)**	(.0002)***	(.0002)***	(.0003)***
Age/10	.0703	-.0224	-.0418	-.0255
	(.0017)***	(.0008)***	(.0012)***	(.0012)***
Age squared/100	-.0076	.0025	.0045	.0027
	(.0002)***	(.0001)***	(.0001)***	(.0002)***
Low education	-.0315	.0042	.0071	.0220
	(.0006)***	(.0003)***	(.0004)***	(.0005)***
High education	.0148	-.0021	-.0036	-.0104
	(.0004)***	(.0002)***	(.0003)***	(.0003)***
Observations	2,112,996	2,112,996	2,112,996	2,112,996
Tenure < 1 year				
	Retention	Quit	Quit2	Layoff
Female	.0112	.0070	-.0000	-.0120
	(.0012)***	(.0005)***	(.0008)	(.0009)***
Age/10	.0444***	-.0243***	-.0390***	-.0019
	(.0052)***	(.0023)***	(.0034)***	(.0039)
Age squared/100	-.0060	.0029	.0045	.0011
	(.0007)***	(.0003)***	(.0005)***	(.0006)*
Low education	-.0571	.0074	.0120	.0421
	(.0019)***	(.0008)***	(.0012)***	(.0015)***
High education	.0412	-.0062	-.0121	-.0276
	(.0015)***	(.0006)***	(.0010)***	(.0010)***
Observations	448,508	448,508	448,508	448,508

Notes. Dependent variable of the retention column: binary variable equals 1 if the worker on a job in month t does not leave the job in the next month. Dependent variable of the remaining columns: binary variable equals 1 if the worker on a job in month t leaves the job in the next month by each route. All regressions include a full set of time (month and years) and region dummies. Weighted standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table 6: Linear Probability Model: Retention, Quits and Layoff - UR effect

	No Controls				With Controls			
	Retention	Quit	Quit2	Layoff	Retention	Quit	Quit2	Layoff
All tenure								
UR	.0184	-.0601	-.0768	.1488	.1158	-.0777	-.1145	.0908
	(.0089)**	(.0041)***	(.0058)***	(.0065)***	(.0090)***	(.0041)***	(.0059)***	(.0065)***
Observations	2,112,996	2,112,996	2,112,996	2,112,996	2,112,996	2,112,996	2,112,996	2,112,996
Tenure < 1 year								
UR	-.4942	-.0851	-.0251	.6446	-.3687	-.1128	-.0740	.5699
	(.0306)***	(.0129)***	(.0199)	(.0235)***	(.0307)***	(.0130)***	(.0201)***	(.0235)***
Observations	448,508	448,508	448,508	448,508	448,508	448,508	448,508	448,508
1 year ≤ Tenure < 3 years								
UR	.1154	-.0694	-.1045	.0874	.1662	-.0901	-.1356	.0698
	(.0204)***	(.0103)***	(.0141)***	(.0140)***	(.0206)***	(.0104)***	(.0143)***	(.0140)***
Observations	410,177	410,177	410,177	410,177	410,177	410,177	410,177	410,177
3 years ≤ Tenure < 7 years								
UR	.1267	-.0584	-.0925	.0456	.1657	-.0711***	-.1103	.0263**
	(.0153)***	(.0076)***	(.0100)***	(.0107)***	(.0155)***	(.0077)***	(.0102)***	(.0108)**
Observations	467,891	467,891	467,891	467,891	467,891	467,891	467,891	467,891
7 years ≤ Tenure < 12 years								
UR	.1242	-.0426	-.0748	.0260	.1526***	-.0481	-.0849	.0056
	(.0154)***	(.0076)***	(.0097)***	(.0104)**	(.0156)***	(.0078)***	(.0100)***	(.0105)
Observations	319,373	319,373	319,373	319,373	319,373	319,373	319,373	319,373

Notes. Dependent variable of the retention columns: binary variable equals 1 if the worker on a job in month t does not leave the job in the next month. Dependent variable of other columns: binary variable equals 1 if the worker on a job in month t leaves the job in the next month by each route. The regression with controls include controls for gender, age, age squared, and dummy variables for education (low and high) and provinces. Weighted standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

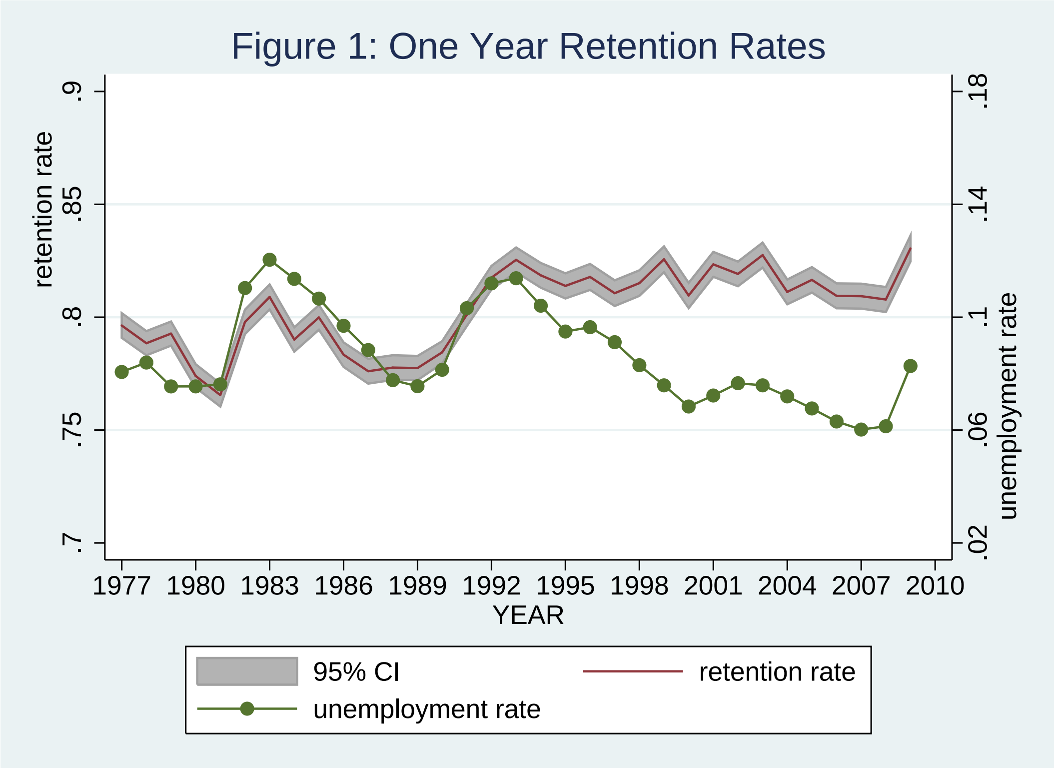
Figure 1: One Year Retention Rates
retention rate
.7
.75
.8
.85
.9
unemployment rate
.02
.06
.1
.14
.18
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
95% CI
retention rate
unemployment rate

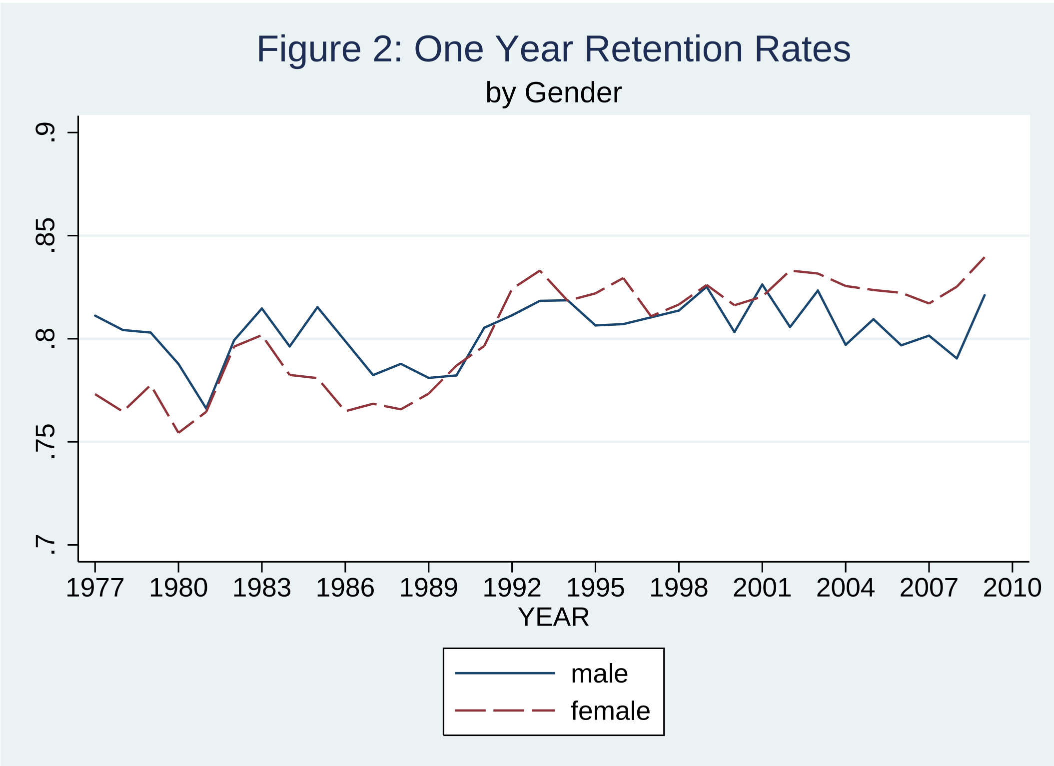
Figure 2: One Year Retention Rates
by Gender
.7
.75
.8
.85
.9
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
male
female

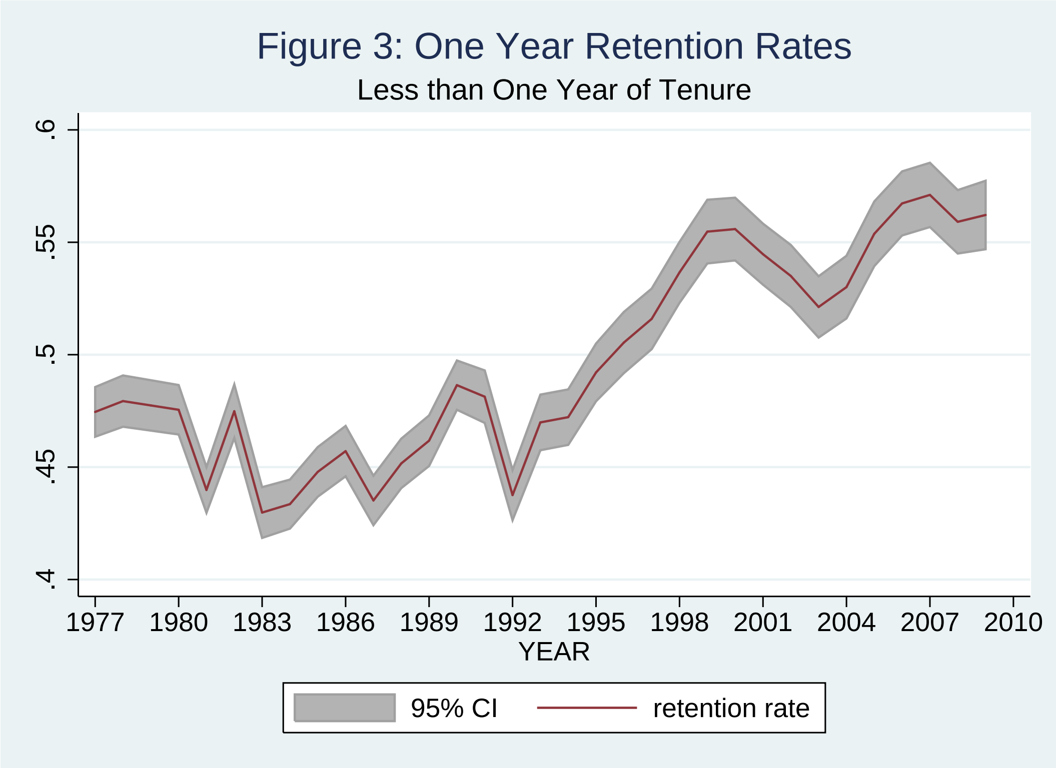
Figure 3: One Year Retention Rates
Less than One Year of Tenure
.4
.45
.5
.55
.6
1977 1980 1983 1986 1989 1992 1995 1998 2001 2004 2007 2010
YEAR
95% CI
retention rate

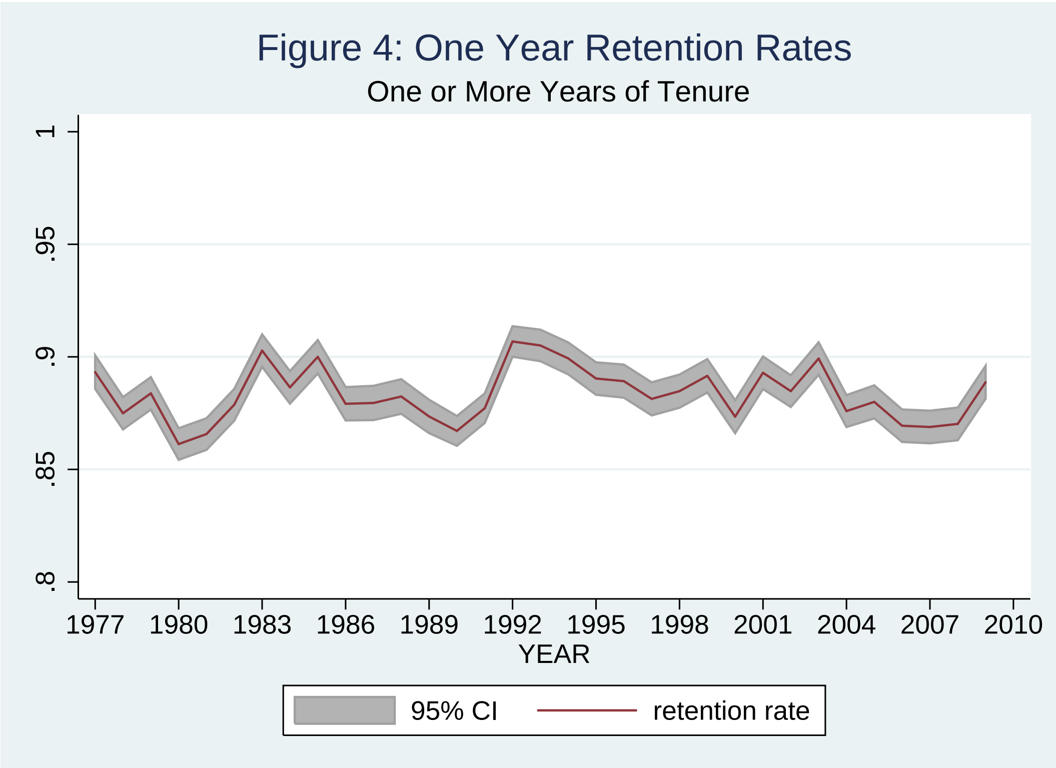
Figure 4: One Year Retention Rates
One or More Years of Tenure
.8
.85
.9
.95
1
1977 1980 1983 1986 1989 1992 1995 1998 2001 2004 2007 2010
YEAR
95% CI
retention rate

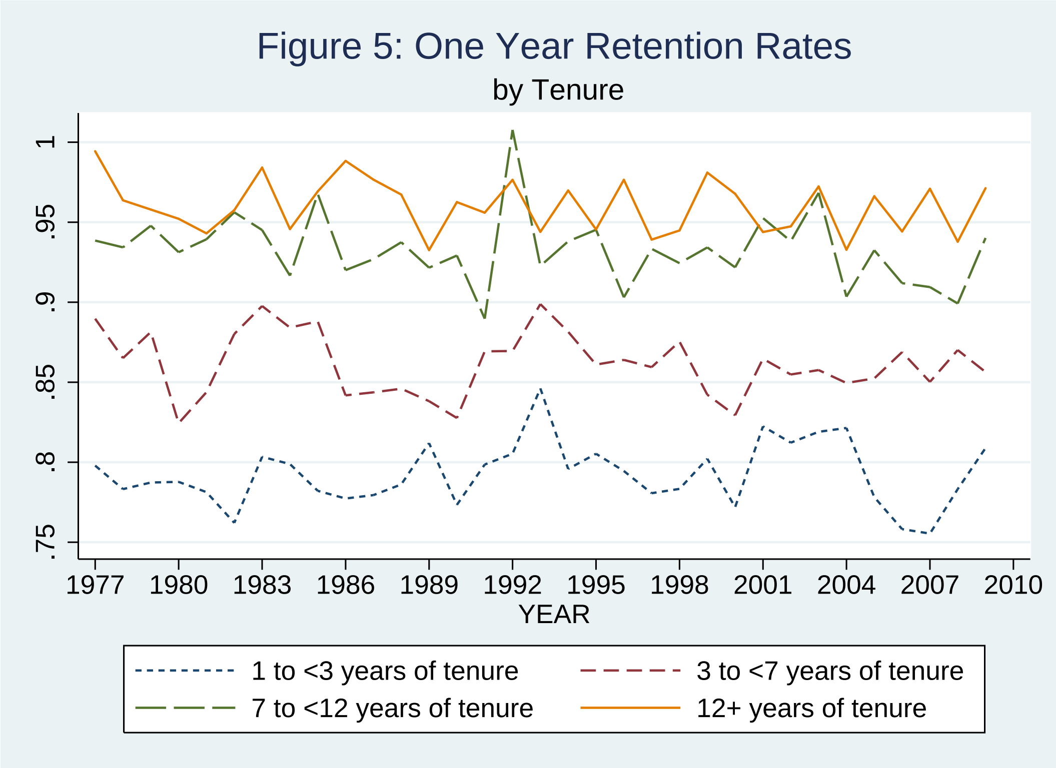

Figure 5: One Year Retention Rates
by Tenure
.75
.8
.85
.9
.95
1
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
1 to <3 years of tenure
3 to <7 years of tenure
7 to <12 years of tenure
12+ years of tenure

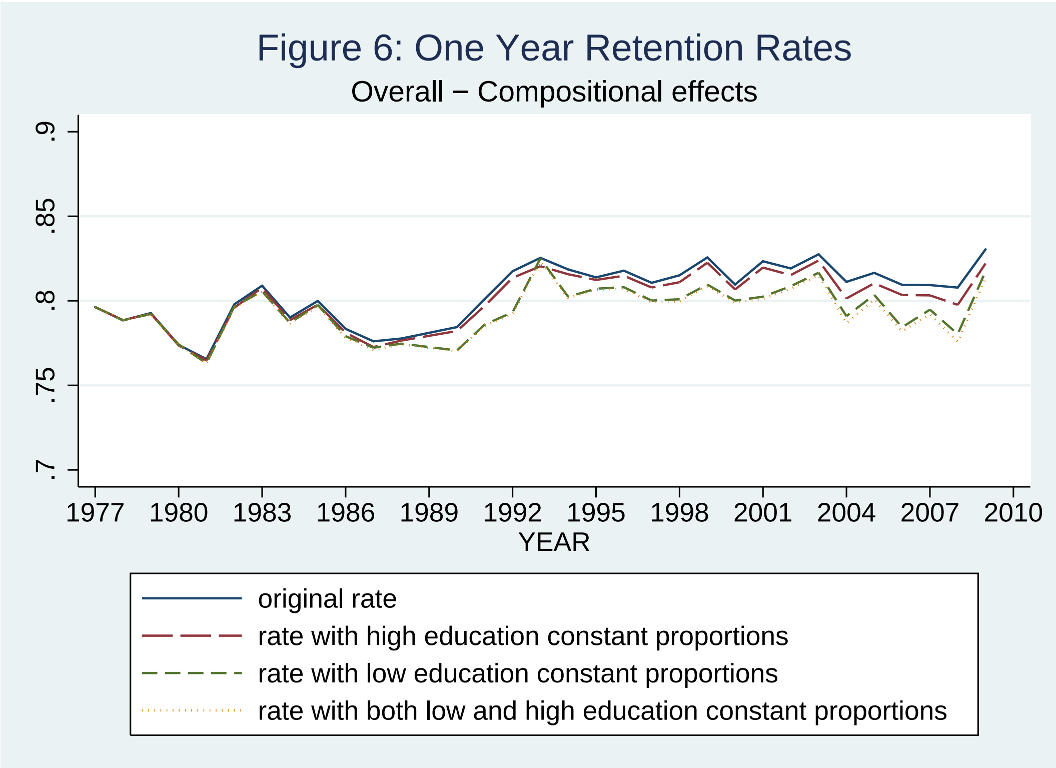

Figure 6: One Year Retention Rates
Overall − Compositional effects
.7
.75
.8
.85
.9
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
original rate
rate with high education constant proportions
rate with low education constant proportions
rate with both low and high education constant proportions

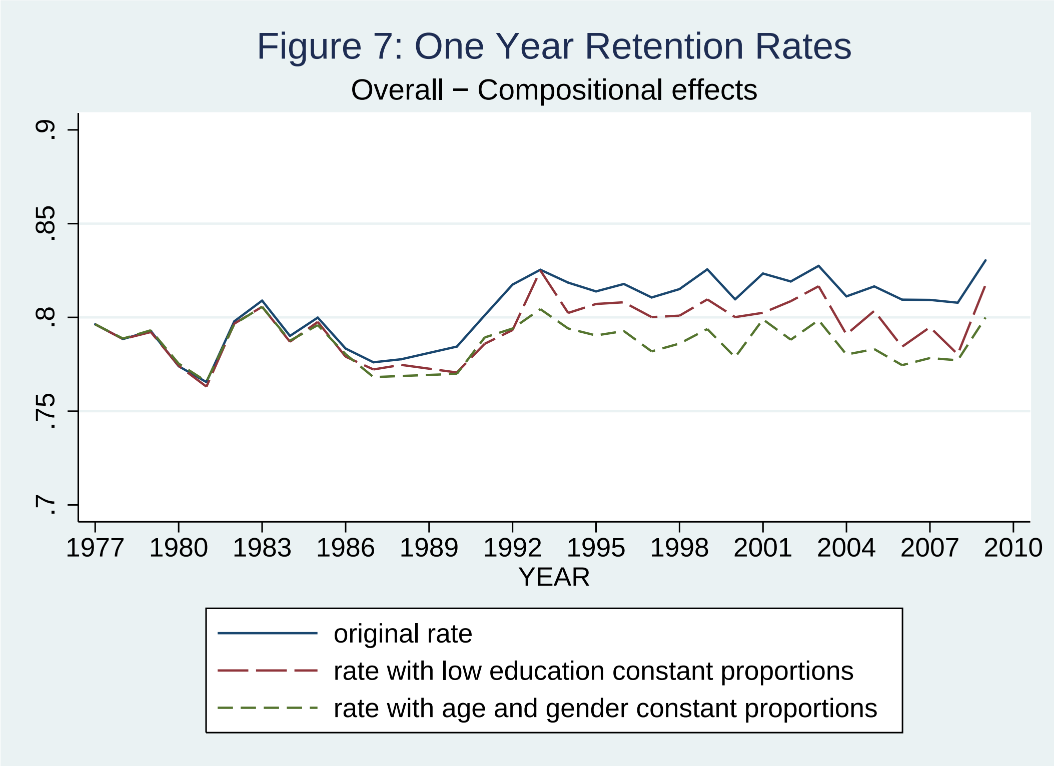

Figure 7: One Year Retention Rates
Overall − Compositional effects
.7
.75
.8
.85
.9
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
original rate
rate with low education constant proportions
rate with age and gender constant proportions

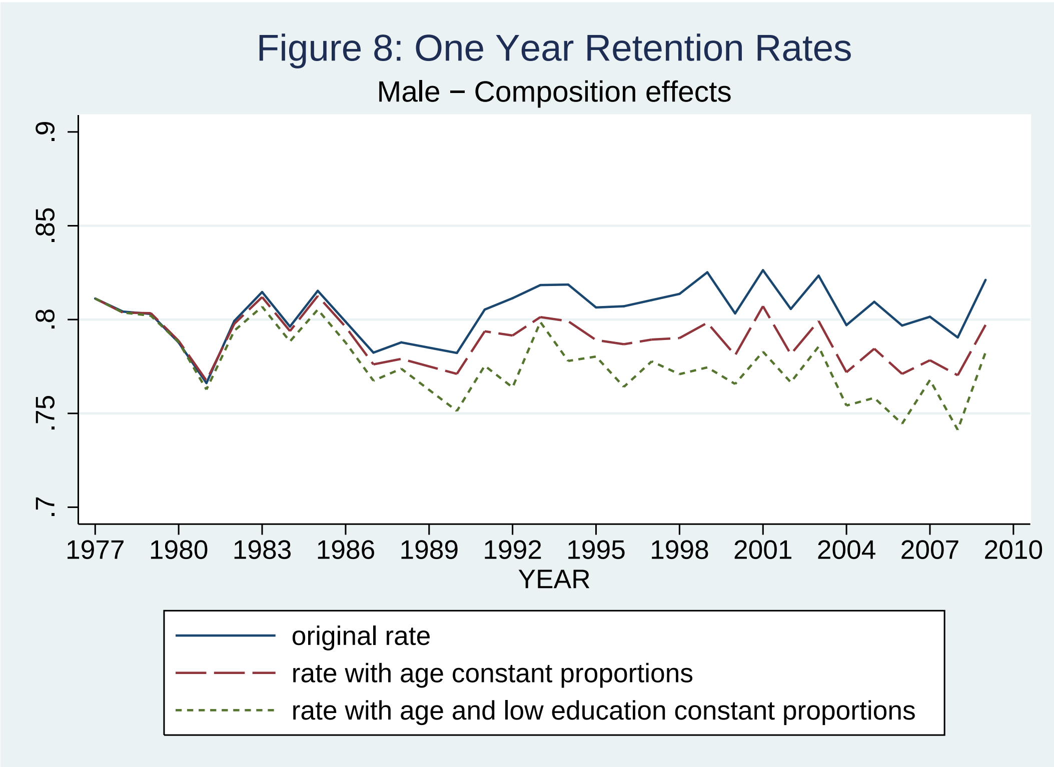

Figure 8: One Year Retention Rates
Male − Composition effects
.7
.75
.8
.85
.9
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
original rate
rate with age constant proportions
rate with age and low education constant proportions

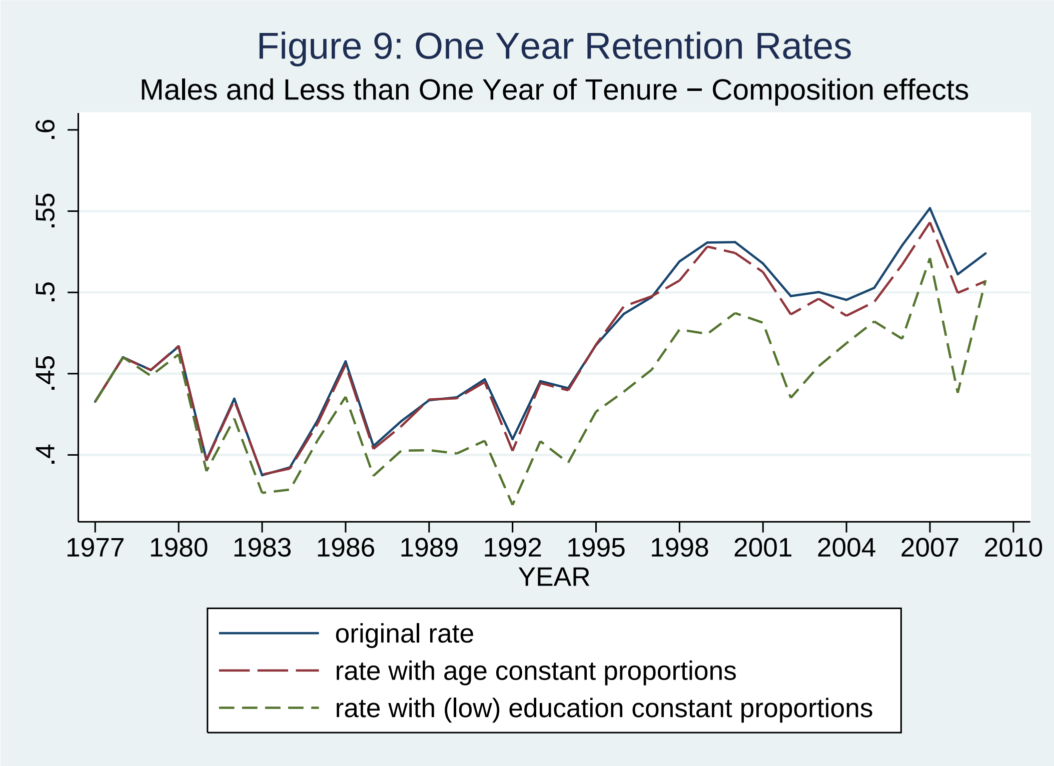
Figure 9: One Year Retention Rates
Males and Less than One Year of Tenure − Composition effects
.4
.45
.5
.55
.6
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
original rate
rate with age constant proportions
rate with (low) education constant proportions

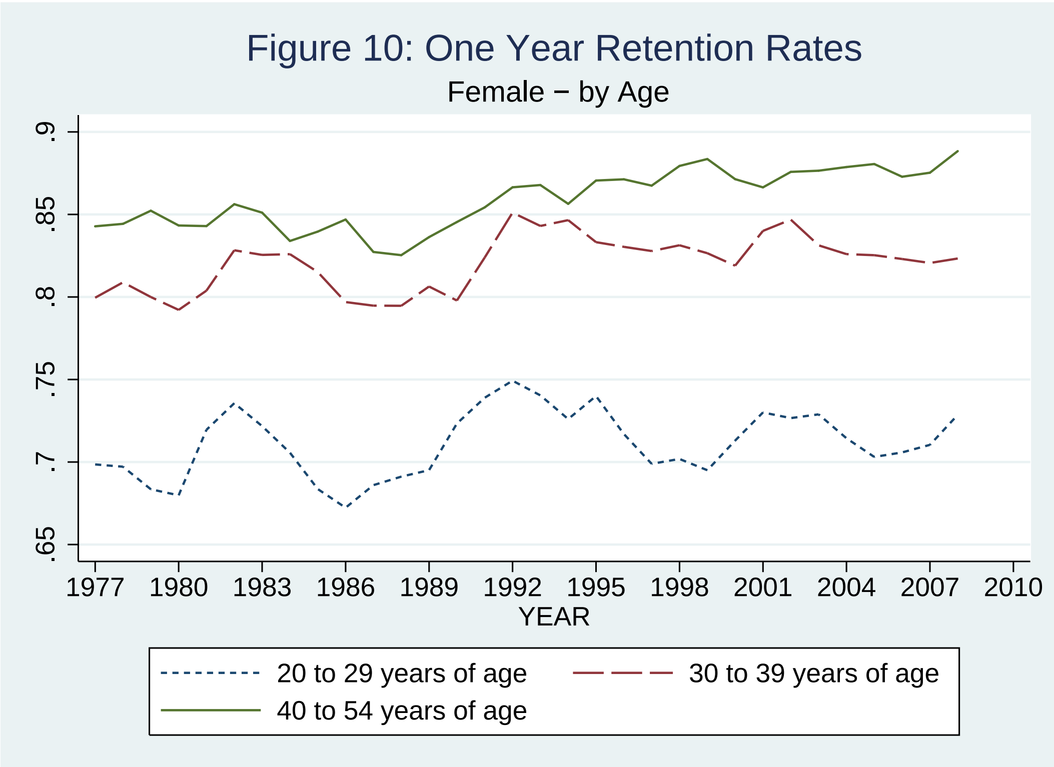
Figure 10: One Year Retention Rates
Female − by Age
.65
.7
.75
.8
.85
.9
1977
1980
1983
1986
1989
1992
1995
1998
2001
2004
2007
2010
YEAR
20 to 29 years of age
30 to 39 years of age
40 to 54 years of age

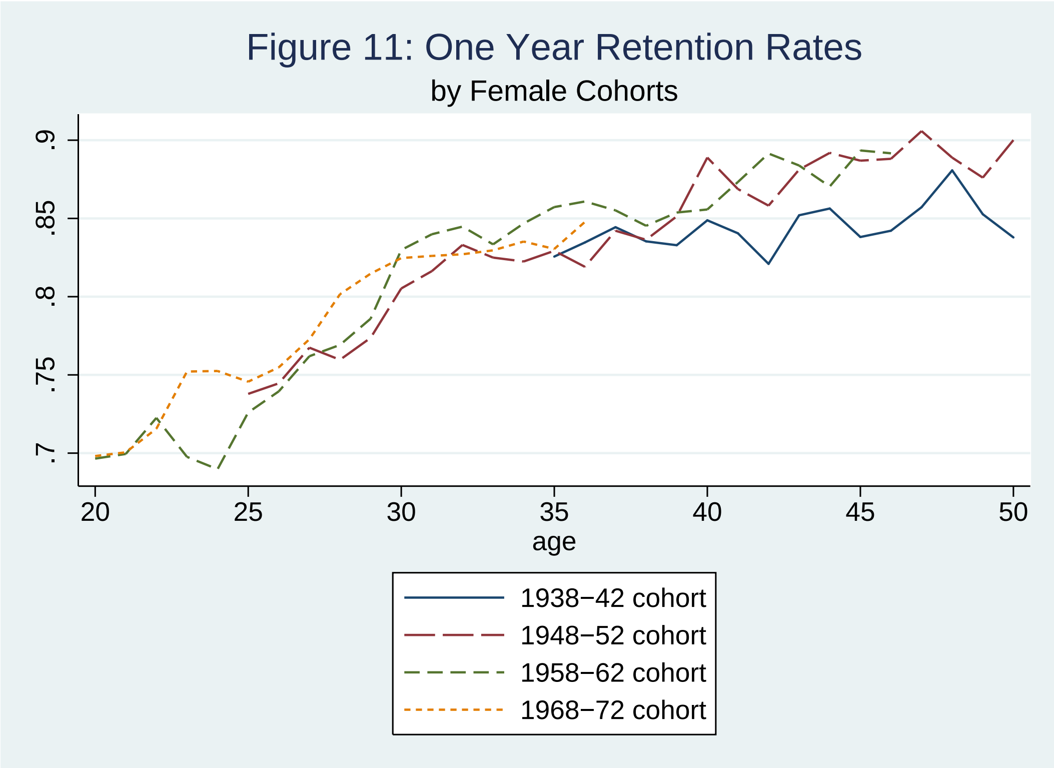
Figure 11: One Year Retention Rates
by Female Cohorts
.9
.85
.8
.75
.7
20
25
30
35
40
45
50
age
1938−42 cohort
1948−52 cohort
1958−62 cohort
1968−72 cohort

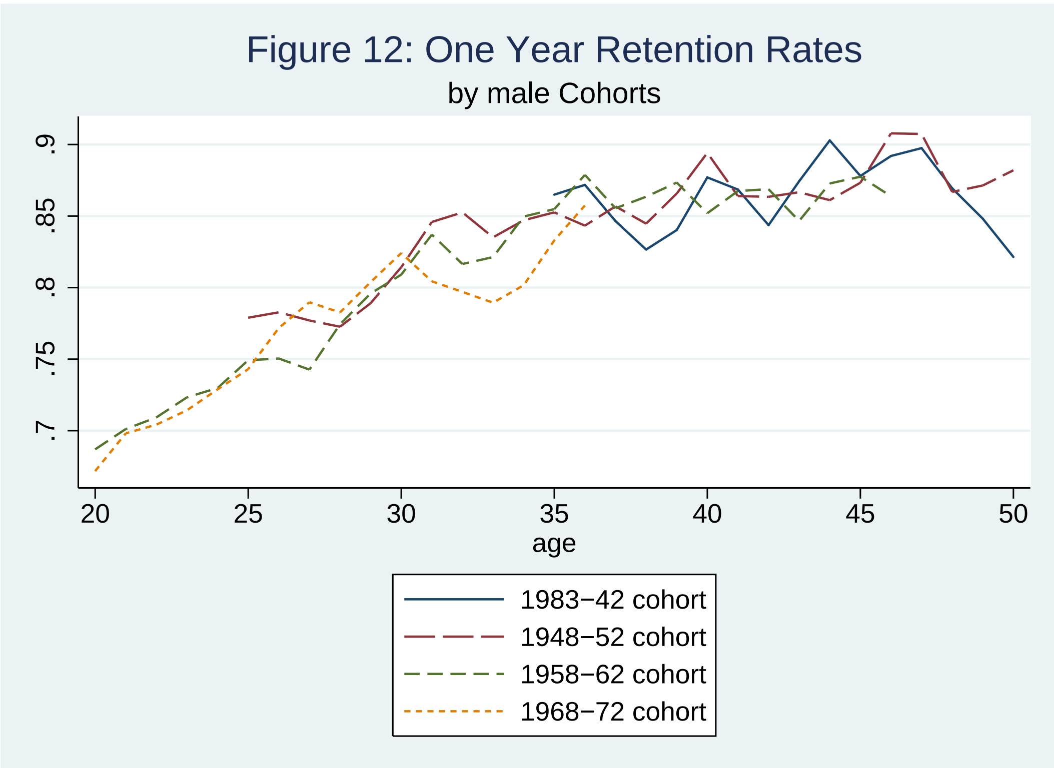
Figure 12: One Year Retention Rates
by male Cohorts
.9
.85
.8
.75
.7
20
25
30
35
40
45
50
age
1983−42 cohort
1948−52 cohort
1958−62 cohort
1968−72 cohort

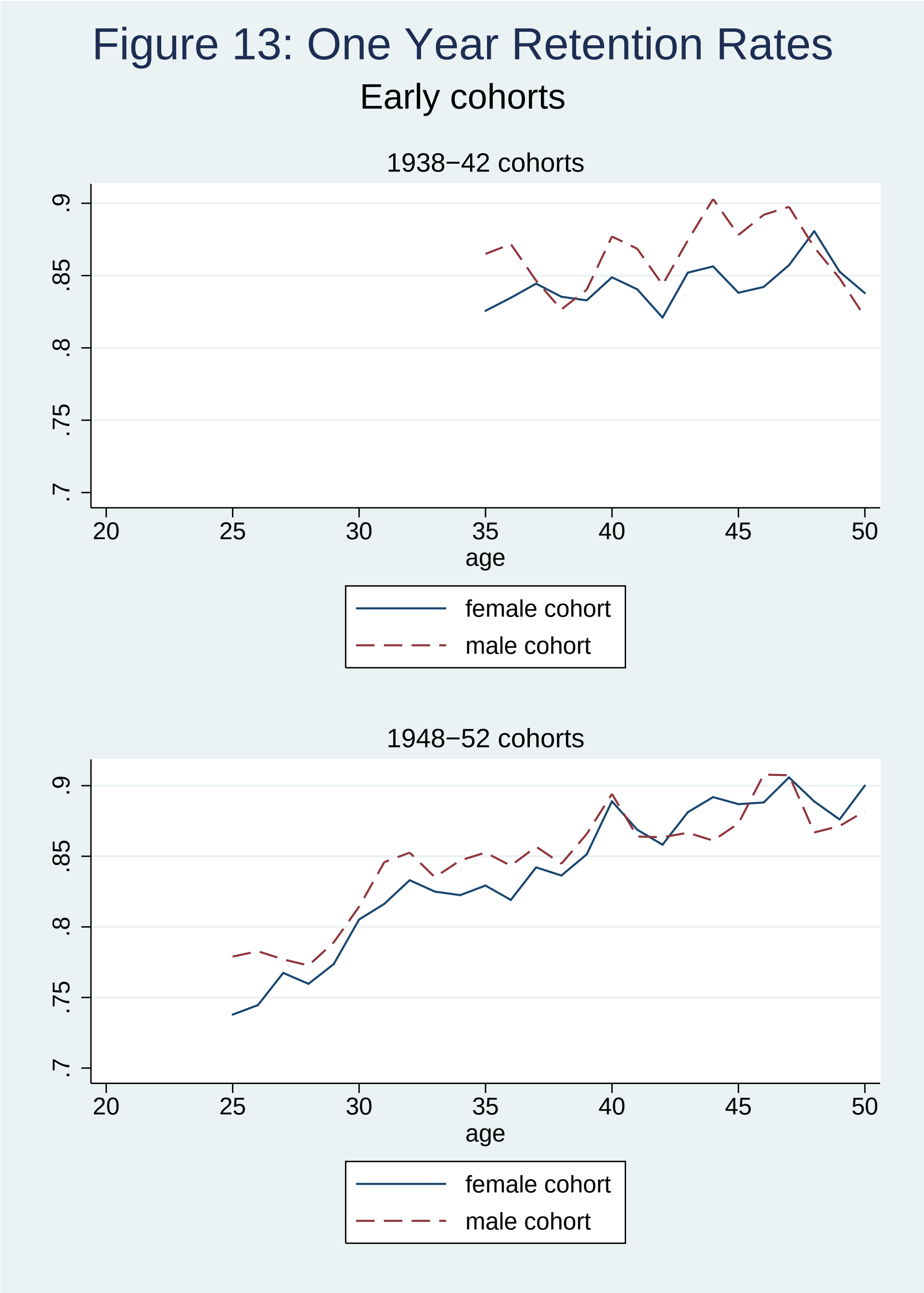

Figure 13: One Year Retention Rates
Early cohorts
1938−42 cohorts
.7
.75
.8
.85
.9
20
25
30
35
40
45
50
age
female cohort
male cohort
1948−52 cohorts
.7
.75
.8
.85
.9
20
25
30
35
40
45
50
age
female cohort
male cohort

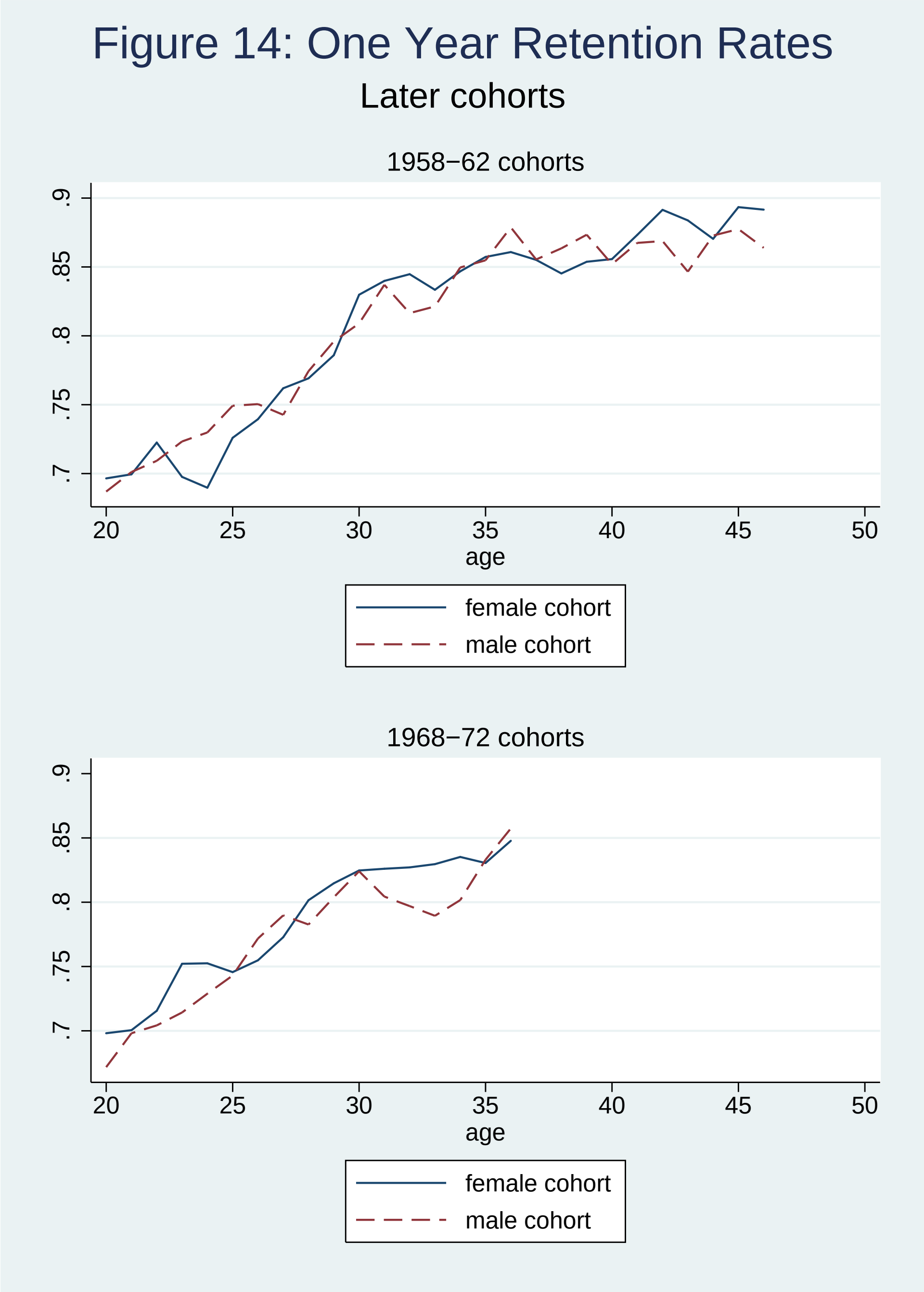
Figure 14: One Year Retention Rates
Later cohorts
1958−62 cohorts
.7
.75
.8
.85
.9
20
25
30
35
40
45
50
age
female cohort
male cohort
1968−72 cohorts
.7
.75
.8
.85
.9
20
25
30
35
40
45
50
age
female cohort
male cohort