CYCLICAL INTERINDUSTRY LABOUR MARKET TRANSITION MECHANISMS: A LOOK AT THE CANADIAN LABOUR MARKET

by
Jimmy Li
5901636

Under the supervision of Dr. Casey Warman

An essay submitted to the Department of Economics
in partial fulfillment of the requirements for
the degree of Master of Arts

Queen's University
Kingston, Ontario, Canada
August 2009

Abstract

Recent literature on the US labour market has focused on a compensating differentials hypothesis where wage cyclicality differentials between job changers and job stayers are explained as compensation for employment risk. This paper sets out to test the robustness of the compensating differentials hypothesis using Canadian data, and provides several extensions to previous studies in the US. Overall, the compensating differentials hypothesis was found to explain Canadian labour market behaviour over the course of the business cycle. However, robustness tests with age, education and income specifications suggest composition bias, where non-middle aged, less educated, and lower income subgroups are exclusively driving the empirical results in support of the compensating differentials hypothesis.

Table of Contents

I. Introduction

The effect of business cycles on the labour market is a valuable relationship to understand. In an economic upturn, previous literature has shown a shift of labour from low wage industries to high wage industries (Reder 1955, Okun 1973). Further evidence has demonstrated that wages are more procyclical for job changers than for job stayers (Okun 1973, Vroman 1977, 1978). This paper attempts to empirically verify a model that will explain the difference in the wage cyclicality between job changers and job stayers.

There are two main streams of thought on this issue. The first returns to the original Okun (1973) model. Here, the argument is that during the business cycle, high wage industries are more procyclical than low wage industries. Thus, during economic upturns, greater labour demand by high wage industries loosens their employment rationing behaviour, thereby allowing job changers to transition to high wage jobs from low wage jobs. During economic downturns the reverse mechanism holds with tighter employment rationing forcing the original job changers to move back into low wage jobs. Hence, wage differentials and employment rationing are the key factors behind the differences in wage cyclicality between job changers and job stayers over the course of the business cycle. A second approach by Barlevy (2001) is based on empirical evidence that job changers experience long unemployment spells after the transition out of high wage industries. A compensating differentials hypothesis was proposed where job changers face higher unemployment risk underlying their position in the high wage industry. Thus, job changers who move to higher wage jobs during economic upturns are merely compensated for future wage losses during economic downturns.

This paper follows the methodology of the second approach by Barlevy (2001) and tests the compensating differentials hypothesis with the Canadian Survey of Labour and Income Dynamics (SLID) as opposed to US data used in the original paper. Further insight will be obtained through the robustness tests under different specifications such as age and education level. The essential relationship to examine is that between employment insurance (EI) benefits and the procyclical behaviour of wages for job changers. A subtle implication of Barlevy's (2001) compensating differentials hypothesis is that if high wages for job changers are to compensate them for future losses of risky employment tenure, then high employment insurance will tend to dampen wage cyclicality among job changers. The opposite conclusion is reached with Okun's (1973) model in which high employment insurance will tend to exaggerate wage cyclicality since EI increases the reservation wage, thereby the bargaining power of job changers. The different conclusions implied by the two hypotheses will allow us to distinguish the dominant mechanism underlying the difference in wage procyclicality between job changers and job stayers.

II. Literature Review

In the past 50 years, much effort has been committed into the research of labour market dynamics during business cycles. Early work by Reder (1955) and Okun (1973) focused on an interesting phenomenon of interindustry transitions. Okun (1973) modeled the shift in labour composition from low paying jobs to high paying jobs during the expansion periods of the business cycle. The model assumed intrinsic qualities of workers by dividing the labour force into quality A and quality B workers. Quality A workers are highly skilled while quality B workers are low skilled. In the absence of business cycle volatilities, the former and the latter would remain in high-wage and low wage jobs, respectively. Further assumptions include perfect competition where marginal productivity is equivalent to wage. Thus, high wage jobs, corresponding to higher productivity sectors, are the main destinations for low wage earners during booms. The opposite case was also hypothesized by Okun. During a downturn, labour shifts from high wage jobs to low wage jobs. For example, quality A workers accept low paying jobs to avoid unemployment.

These patterns of interindustry transition bring us to the cyclicality of these high-paying jobs. Building on the work by Okun, Vroman (1977, 1978) has shown that high-paying sectors are more sensitive to the business cycle than low-paying sectors. Quality B workers account for the majority of the shifts into high wage jobs during booms. Furthermore, these workers also account for the majority of the exits from high wage jobs during downturns. Using National Longitudinal Survey of Young Men with data starting in 1966, Bils (1985) has shown that wage cyclicality is present only amongst job changers. Similarly, with Panel Study of Income Dynamics data, Shin (1994) was able to confirm that wages were more cyclical for job changers than for job stayers over the course of the business cycle.

These labour market shifts lead naturally to the question of what are the economic motivators for these labour composition shifts during the business cycle. Okun (1973) explains this simply with wage differentials and labour demand. High wage sectors have higher productivity and higher procyclical employment behaviour compared to low wage sectors. Thus, during an economic upturn, the demand for labour from these sectors exceeds the supply leading to the incorporation of some quality B workers into the workforce. The argument may be too simple because it cannot resolve a fundamental question: why is there a difference in the wage cyclicality between job changers and job stayers? A modern approach to this problem was proposed by Beaudry and DiNardo (1991), who introduced an implicit contract model to explain the wage cyclicality differences. Job stayers are protected by contracts with their current employers and thus are shielded partially or completely against the ups and downs of the business cycle. On the other hand, job changers have minimal contractual agreements with their new employers and have little bargaining power with the wage that is offered to them. This wage is highly dependent on prevailing economic conditions at the time of the offer.

The only other modern attempt at understanding the difference in the wage cyclicality between job changers and job stayers was made by Barlevy (2001). A compensating differentials hypothesis was proposed in which job changers face higher unemployment risk underlying their position in high wage industries. Thus, job changers who move to higher wage jobs during economic upturns are merely compensated for future wage losses during economic downturns. This leads naturally to a model where the true gains from the transition to higher wage jobs during economic booms may not be realized.

A small quantity of older literature supports these claims. Abowd and Ashenfelter (1981) modelled competitive equilibrium wage with endogenous anticipated unemployment and unemployment risk. The certainty equivalent compensation was found to be proportional to the expected unemployment rate squared. Furthermore, the risk compensation was determined to be proportional to the coefficient of unemployment variation. Depending on the industry, a compensation differentials running from less than 1% to more than 14% of wages was found. Thus, there is significant evidence that workers' wages and unemployment risk are correlated ceteris paribus. Similarly, Murphy and Topel (1987) established that some of the interindustry wage gaps could be explained by interindustry unemployment risk differences. Furthermore, Topel (1984) found that unemployment insurance has a strong and significant effect on the magnitude of the compensating wage premium. Thus, the final link is to extend these models in a similar effect to wage cyclicality in lieu of wages.

The data used to substantiate the Barlevy (2001), Abowd and Ashenfelter (1981), Murphy and Topel (1987) and Topel (1984) conclusions includes the Panel Study of

Income Dynamics (PSID) and the National Longitudinal Survey of Youth (NLSY), both of which were based on the US population. Despite the various robustness tests and theoretical models used, the compensating differentials models have not been tested using non-US data, so it is not known if the results hold outside of the US. Therefore, this paper attempts to apply Barlevy`s (2001) empirical model to the Canadian Survey of Labour and Income Dynamics (SLID). Moreover, composition bias is a factor that has not been explored in the previous literature. That is, certain components of the sample group may be driving the results of the entire sample. Segmentation of the data by specifications such as age, education and income groups will help to identify the presence or absence of such bias.

III. Methodology and Data

The fundamental methodology of this study follows that of Barlevy (2001) with additional specifications such as fracturing the sample into different age, education and income groups. The main prediction is based on the implications of the compensating differentials hypothesis. That is, areas with higher EI levels will exhibit lower wage cyclicality for job changers compared to areas with lower EI levels. The fixed effects model is:

$$\Delta \ln W_{it} = \beta_1 + \beta_2 t + (\alpha_1 + \alpha_2 EI_{it})\Delta U_{it} + \beta_4 X_{it} + v_{it},$$

where W_{it} is person i's real wages at time t. This is calculated by dividing the nominal wage by the CPI deflator using 1999 as the base year. SLID reports both composite hourly wage and total salaries and wages. The former is considered in this model while the latter is not used. The results using the latter measure underestimate the total wage gain for job changers because this data group tends to have periods of unemployment. On the other hand, the composite hourly wage is the average hourly wage rate weighted by the hours worked in each job. Thus, this measure is ideal because it does not take into account periods of unemployment associated with job changers. U_{it} is the unemployment rate in the EI region where person i resides at time t. This is averaged from the monthly unemployment rates calculated by Human Resources Social Development Canada (HRSDC). EI_{it} is the employment insurance benefits level for person i at time t. Three measures are considered for the regression: the maximum number of weeks of EI that the respondents are eligible for, the minimum number of weeks of EI that respondents are eligible for and the total number of work hours required for a respondent to be eligible for employment insurance. The first two measures positively correspond to higher EI benefits. The data for the last measure is inversed to reflect its negative relationship to higher EI benefits. That is, a lower number of work hours required for a respondent to be eligible for employment insurance indicates a higher level of EI benefits and vice versa. X_{it} is the work experience of person i at time t. Finally, v_{it} is the error term.

The immediate problem with this regression is that the explanatory variable $\beta_3 = \alpha_1 + \alpha_2 EI_{it} + \varepsilon_{it}$ is serially correlated with v_{it}. That is, the error terms of different individuals have analogous time effects. Regression by ordinary least squares would lead to biased and inconsistent estimates. Thus, as proposed by Solon, Barsky, and Parker (1994), a two-stage least squares provides the best estimates. In the first stage, ΔlnW_{it} is regressed on X_{it} and EI region state-year dummy variables. The coefficients from these state-year dummy variables are saved and regressed on ΔU_{it}, $EI_{it} \cdot \Delta U_{it}$ and a time trend in the second stage. This step is weighted by the populations in each respective employment insurance region. The primary aim is to observe the interaction term between EI_{it} and ΔU_{it}, where a positive relationship would support the compensating differentials hypothesis.

Finally, it is important to identify the job changers who are the subject of the above regression. SLID lists a variable that incorporates the reason for job separation. Polsky (1999) advocates that any individual that answers this survey question except for reasons of internal promotion should be considered a job changer. This paper will use this measure to identify job changers. The job changer sample focuses mainly on male individuals of working age (16-65). The reasoning is outlined by Solon, Barsky, and Parker (1994). Women have very low participation rates in EI programs compared to men. Furthermore, women form a smaller proportion of the labour force at high wage procyclical industries such as manufacturing and construction. Both of these realities minimize the effect of EI benefits on female wage cyclicality. Despite this, a small digression will be explored with a regression of a female only and gender combined sample.

Data is taken from the Canadian Survey of Labour and Income Dynamics (SLID). This data is appended from five sets of overlapping longitudinal cohorts with years: 1992-1998, 1995-2001, 1998-2004, 2001-2007, and 2004-2007. In each cohort, individuals are tracked and interviewed for all six years (except for the last cycle which only has four years of data). Basic summary statistics are presented in Table 1. Around 13 percent of the sample changes jobs. Job changers are individuals who change jobs at least once. Job changers are generally younger, less well paid, less educated, less experienced, and less involved in unions than people who do not change jobs.

Table 1: Summary Statistic

	Job Changers	Job Stayers
% of sample	13.2	86.8
Age	33.3	38.7
Real Composite Hourly Wage (1999$)	12.9	17.24
Years of Schooling	16.7	17.5
Years of Work Experience	13.03	17.99
%Union	0.241	0.368
%Durable Manufacturing	0.04	0.04

Notes: The sample is restricted to males between the ages of 16 – 65. Job changers are defined as any individual who has answered the SLID survey question regarding the reason for job separation with any answer other than internal promotion. Other than job changers, all remaining individuals of the sample are defined as job stayers.

IV. Results

The results of the fixed effects model are outlined in Tables 2 to 9 with p-values in parenthesis. To varying degrees of statistical significance, the interaction terms EI • ΔUit are generally positive throughout. Intuitively, one must remember that unemployment and wage have a negative relationship. Thus, a positive interaction term implies that in areas of higher EI levels, wage cyclicality for job changers decreases. These results are consistent with that predicted by the compensating differentials hypothesis. Furthermore, there is no evidence of the interindustry wage differentials model as no interaction term coefficient is negative and statistically significant in any of the regressions in this study.

Table 2 : Job changer wage cyclicality

	Maximum EI	Minimum EI	Qualifying EI
Constant	1.9767 (0.000)	1.9871 (0.000)	1.9938 (0.000)
ΔUit	-.0307 (0.032)	-.0127 (0.045)	-.0169 (0.023)
EI • ΔUit (/100)	.0563 (0.037)	.0292 (0.057)	.0605 (0.029)
F-test value	29.09	28.86	29.22
R^2	0.19	0.19	0.19
Years	1997-2007	1997-2007	1997-2007
N	16,249	16,249	16,249

Notes: In the first stage regression, work experience and region-time dummy variables are regressed on Δln W and weighted by sampling weights. W is the real composite hourly wage (indexed by CPI to 1999$), the regions are employment insurance regions and the time is by year. In the second stage, ΔU, EI • ΔUit and a time trend are regressed on the dummy coefficients of the first stage and weighted by the population in each region to produce the results shown. ΔU_{it} is the unemployment rate in each employment insurance region t in year i. EI is the employment insurance level in each employment insurance region. Three regressions are performed with three different measures of EI levels. Maximum EI and minimum EI is the maximum and minimum number of weeks of EI a respondent is eligible for, respectively. Qualifying EI is the inverse of the number of weeks of work a respondent has to perform before becoming eligible to receive EI. N is the number of observations in the first stage regression. Job changer is defined in Table 1. P-values are in parenthesis.

Table 2 shows a positive interaction term between EI_{it} and ΔU_{it} which is statistically significant at the five percent level for composite real hourly wage. The interaction term estimates indicate that wage cyclicality will increase by 0.0563%, 0.0292% and 0.0605% for every 100 hour increase in maximum, minimum and qualifying EI levels, respectively. The results may appear small in magnitude; however, they are in the opposite direction of the Okun (1973) model of interindustry wage differentials and are statistically significant, and in favour of the Barlevy (2001) compensating differentials hypothesis.

Overall, the estimates show greater significance with maximum EI and qualifying EI levels. This is not surprising, as research by Andersen and Meyer (1993) has demonstrated that approximately one out of three unemployment insurance recipients take the maximum benefit offered. Overall, the compensating differential hypothesis is demonstrated more consistently and significantly using maximum EI level as a measure.

Table 3: Job changer wage cyclicality controlling for changes in union and manufacturing composition

	Maximum EI	Minimum EI	Qualifying EI
Constant	2.1164	2.1084	2.1106
	(0.000)	(0.000)	(0.000)
ΔUit	-.0409	-.0177	-.0201
	(0.041)	(0.051)	(0.059)
EI • ΔUit (/100)	.0716	.0372	.0655
	(0.053)	(0.083)	(0.091)
Δ union	.0886	.1096	.1012
	(0.489)	(0.392)	(0.430)
Δ manufacturing	-.0560	-.0574	-.0519
	(0.624)	(0.616)	(0.651)
F-test value	12.62	12.40	12.36
R^2	0.43	0.42	0.42
Years	1997-2007	1997-2007	1997-2007
N	16,249	16,249	16,249

Notes: Δ union is equal to +1 if a job changer transitions from a non-union position to a union or collective bargaining position. Δ union is equal to -1 if a job changer transitions from a union or collective bargaining position to a non-union position. A value of 0 for Δ union indicates no change in unionization. Δ manufacturing is equal to +1 if a job changer transitions into durable manufacturing from other industries. Δ manufacturing is equal to -1 if a job changer transitions out of durable manufacturing into other industries. A value of 0 for Δ manufacturing indicates no transition into or out of manufacturing. Additional required definitions and notes are in table 2. P-values are in parenthesis.

While the results in table 2 are of significance, there exist many exogenous sources of error. One issue is the implicit assumption of zero or little correlation between regional heterogeneity and EI benefit levels. In his study of US data, Barlevy (2001) argues that the extent of unionization and the industrial composition of the region should be controlled for. The former affects wage distributions and thus, the wage cyclicality of job changers. Moreover, the latter is important because according to the interindustry wage differentials hypothesis, cyclical industries (such as durable manufacturing as pointed out by Barlevy (2001)) have a strong incentive to move to areas of low EI benefits. This is because EI levels act as a reservation wages. Lower EI levels results in lower reservation wages thereby giving firms more bargaining power with workers in wage negotiations. Thus, if a correlation between regional heterogeneity and EI benefit levels does exist, a positive relationship between EI benefits and wage cyclicality for job changers would naturally arise.

Table 3 illustrates the results of the empirical model with controls for changes in union and manufacturing composition. In the regression, Δ union is equal to +1 if a job changer transitions from a non-union position to a union or collective bargaining position. Conversely, Δ union is equal to -1 if a job changer transitions from a union or collective bargaining position to a non-union position. A value of 0 for Δ union indicates no change in unionization. A similar system is used for Δ manufacturing, where +1 is equivalent to a shift into a durable manufacturing position, -1 is equivalent to a shift out of a durable manufacturing and 0 signals no job movement between durable manufacturing and all other industries. The data for Δ manufacturing was limited to 1997-1999, and thus is not a complete control. This may explain the greater level of statistical significance for Δ union relative to Δ manufacturing.

Table 3 shows all the positive interaction terms are statistically significant within the ten percent level for composite real hourly wage. The interaction term estimates indicate that wage cyclicality will increase by 0.072%, 0.037% and 0.066% for every 100 hour increase in maximum, minimum and qualifying EI levels, respectively. These results are consistent in direction and magnitude to the results obtained in the basic model without controls (Table 2). The loss of significance in the magnitude of approximately five percent suggests there is some level of correlation between regional heterogeneity and EI levels. Overall, the essential outcome is that the compensating differential hypothesis is supported strongly with the Canadian SLID data.

Table 4: Job stayer wage cyclicality

	Maximum EI	Minimum EI	Qualifying EI
Constant	2.1921	2.1963	2.1935
	(0.000)	(0.000)	(0.000)
ΔUit	-.0013	-.0006	-.0015
	(0.906)	(0.729)	(0.792)
EI • ΔUit (/100)	.0006	-.0363	.0020
	(0.979)	(0.802)	(0.928)
F-test value	60.95	60.97	60.95
R^2	0.33	0.33	0.33
Years	1997-2007	1997-2007	1997-2007
N	106,684	106,684	106,684

Notes: Additional required definitions and notes are in table 2. P-values are in parenthesis.

Other factors that have the potential of driving the wage cyclicality results outlined in table 2 include multiple forms of general regional heterogeneity. For example, regional heterogeneities such as oil price shocks may drive the wages cyclicality patterns in the employment insurance regions of Alberta. Likewise, Barlevy (2001) suggests that price level difference between regions may account for wage cyclicality patterns. This problem cannot be eliminated easily because the wages and prices for each region are adjusted by a national price index instead of a corresponding regional price index.

Table 4 shows the results for job stayers instead of job changers. There is no consistent direction, magnitude or significance across all three measures of EI levels. These results are encouraging because the jobs stayers act as a control group in principle. Any significant exogenous regional-specific factors that account for wage cyclicality in job changers will affect the job stayers as well. Thus, the absence of significant results provides evidence against exogenous factors not accounted for in the job changer empirical model.

Table 5: Job changer wage cyclicality with data from both sexes

	Maximum EI	Minimum EI	Qualifying EI
Constant	1.8829 (0.000)	1.9019 (0.000)	1.9030 (0.000)
ΔUit	-.0121 (0.291)	-.0087 (0.086)	-.0106 (0.075)
EI • ΔUit (/100)	.0220 (0.309)	.0204 (0.097)	.0382 (0.084)
F-test value	42.12	42.69	42.77
R^2	0.25	0.25	0.25
Years	1997-2007	1997-2007	1997-2007
N	32,069	32,069	32,069

Notes: The sample combines the job changers of both genders. Additional required definitions and notes are in table 2. P-values are in parenthesis.

The exclusion of women from the sample group was explained previously as a consequence of low participation rates in EI programs and cyclical industries such as durable manufacturing. However, it would be interesting to discover the consistency of female labour force behaviour with the compensating differential hypothesis. Table 5 presents the results of empirical model with both sexes. In comparison with the results of a male only model (Table 2), the magnitude and direction of the interaction term EI • ΔUit are consistent. However, the statistical significance of the results are depressed with only minimum EI and qualifying EI being statistically significant at least at the ten percent level. Thus, as a complete representation of the Canadian labour market, the compensating differential hypothesis still holds within reasonable extent.

Table 6: Job changer wage cyclicality with data from female sample

	Maximum EI	Minimum EI	Qualifying EI
Constant	1.8221 (0.000)	1.8408 (0.000)	1.8371 (0.000)
ΔUit	.0036 (0.794)	-.0039 (0.529)	-.0034 (0.634)
EI • ΔUit (/100)	-.0079 (0.761)	.0081 (0.583)	.0106 (0.689)
F-test value	29.79	29.86	29.81
R^2	0.19	0.19	0.19
Years	1997-2007	1997-2007	1997-2007
N	15,820	15,820	15,820

Notes: The sample consists only of job changers of the female gender. Additional required definitions and notes are in table 2. P-values are in parenthesis.

Despite the statistically significant results obtained in a sample with only male individuals (Table 2), the existence of the compensating differential hypothesis in a female only sample is still in doubt. The results with a female only sample (Table 6) upholds these doubts as the direction and magnitude of the interaction terms are not consistent. However, the results are not statistically significant at conventional levels (with P-values greater than 0.50). These results endorse studies by Solon, Barsky, and Parker (1994) on US data which conclude that cyclical labour demand shifts over the course of the business cycle are not gender neutral. This has major policy implication for Canadian macroeconomic labour policy. Namely, heterogeneous wage cyclicality between men and women necessitates disparate policy measurements and actions plans.

On the other hand, it may be argued that the statistically insignificant results for a female only sample may be driven by the original reason for their exclusion in previous studies, namely the low participation rates in EI programs and cyclical industries such as manufacturing. By the nature of the two-stage regression used in this study, specifying a first-stage regression with a sample of only women who have participated in EI or durable manufacturing industries results in a small sample size in the second stage regression. Thus, the relevance of the compensating differentials hypothesis on a female sample requires further study with a different methodology, ideally one with a single stage regression.

Along the same line of reasoning, composition bias may exist in the original male only job changer sample. That is, a certain subsample in the male sample may be driving the results of the whole group. Thus, regressions with segmented samples by age (Table 7), years of schooling (Table 8) and income (Table 9) are computed.

For the age group 15-26, the direction and magnitude of the interaction term coefficient supports the compensating differentials hypothesis although only the maximum EI measure are statistically significant at the 10 percent level. The next two age groups 26-35 and 36-45 derive the most surprising results with inconsistent direction and magnitude. Only one in six coefficients is statistically significant at the five percent level. Finally, the last two age groups 46-55 and 56-65 again support the compensating differentials hypothesis with positive interaction terms statistically significant at least at the ten percent level for four out of six EI measures. Thus, the age separation regression affirms composition bias. The compensating differentials hypothesis holds only for the younger and older segments of the Canadian labour force, whereas those in the middle age groups 26-45 show little conformity to the model. The younger and older segments are components of the population driving the results displayed in Table 2 for the general population. The results suggest segmented labour market behaviour over the life cycle in regards to the compensating differentials hypothesis. This is an interesting topic which has not been explored extensively in economic literature and thus requires further study.

Table 7: Job changer wage cyclicality with age group separation

	Maximum EI	Minimum EI	Qualifying EI
	AGE 16 – 25		
Constant	1.8402 (0.000)	1.8054 (0.000)	1.8371 (0.000)
ΔUit	-.0700 (0.041)	-.0171 (0.265)	-.0260 (0.145)
EI • ΔUit (/100)	.1210 (0.059)	.0275 (0.445)	.0756 (0.251)
F-test value	6.88	6.08	6.28
R^2	0.07	0.06	0.06
Years	1997-2007	1997-2007	1997-2007
N	7061	7061	7061
	AGE 26 – 35		
Constant	1.9279 (0.000)	1.7973 (0.000)	1.8942 (0.000)
ΔUit	-.0818	.0079	-.0180

	(0.131)	(0.743)	(0.525)
EI • ΔUit (/100)	.1615	-.0108	.0807
	(0.114)	(0.852)	(0.441)
F-test value	7.49	6.80	6.95
R^2	0.11	0.10	0.10
Years	1997-2007	1997-2007	1997-2007
N	2387	2387	2387
	AGE 36 – 45		
Constant	2.1983	2.1836	2.1667
	(0.000)	(0.000)	(0.000)
ΔUit	.0021	.0108	.0161
	(0.974)	(0.693)	(0.621)
EI • ΔUit (/100)	.0136	-.0036	-.0258
	(0.913)	(0.957)	(0.832)
F-test value	3.54	3.54	3.55
R^2	0.05	0.05	0.05
Years	1997-2007	1997-2007	1997-2007
N	2477	2477	2477
	AGE 46-55		
Constant	2.2000	2.2713	2.2817
	(0.000)	(0.000)	(0.000)
ΔUit	-.1037	-.0497	-.0612
	(0.099)	(0.071)	(0.060)
EI • ΔUit (/100)	.2001	.1322	.2439
	(0.090)	(0.052)	(0.046)
F-test value	3.44	3.68	3.73
R^2	0.05	0.05	0.05
Years	1997-2007	1997-2007	1997-2007
N	2487	2487	2487
	AGE 56-65		
Constant	3.1794	3.1367	3.1617
	(0.000)	(0.000)	(0.000)
ΔUit	-.1812	-.0590	-.0725
	(0.028)	(0.103)	(0.086)
EI • ΔUit (/100)	.3124	.1132	.2207
	(0.042)	(0.206)	(0.161)
F-test value	3.11	2.45	2.54
R^2	0.06	0.05	0.05
Years	1997-2007	1997-2007	1997-2007
N	1719	1719	1719

Notes: The data was segmented into age groups. Additional required definitions and notes are in table 2. P-values are in parenthesis.

Table 8: Job changer wage cyclicality with years of schooling separation

	Maximum EI	Minimum EI	Qualifying EI
	Elementary Education		
Constant	2.6140 (0.000)	2.3691 (0.000)	2.4674 (0.000)
ΔUit	-.2647 (0.006)	-.0338 (0.383)	-.0646 (0.170)
EI • ΔUit (/100)	.4948 (0.005)	.0850 (0.368)	.2449 (0.162)
F-test value	2.47	0.62	0.92
R^2	0.10	0.03	0.04
Years	1997-2007	1997-2007	1997-2007
N	454	454	454
	Secondary Education		
Constant	2.0054 (0.000)	1.9840 (0.000)	1.9954 (0.000)
ΔUit	-.0662 (0.026)	-.0148 (0.265)	-.0210 (0.175)
EI • ΔUit (/100)	.1265 (0.024)	.0382 (0.234)	.0822 (0.155)
F-test value	7.91	6.93	7.09
R^2	0.06	0.06	0.06
Years	1997-2007	1997-2007	1997-2007
N	7268	7268	7268
	Post-Secondary Education		
Constant	2.0203 (0.000)	2.0094 (0.000)	2.0507 (0.000)
ΔUit	-.0086 (0.838)	.0011 (0.952)	-.0111 (0.598)
EI • ΔUit (/100)	.0196 (0.805)	.0017 (0.970)	.0489 (0.535)
F-test value	4.44	4.43	4.53
R^2	0.04	0.04	0.04
Years	1997-2007	1997-2007	1997-2007
N	4887	4887	4887
	Graduate Education		
Constant	2.3717 (0.000)	2.4786 (0.000)	2.5495 (0.000)
ΔUit	-.0608 (0.453)	-.0434 (0.226)	-.0721 (0.089)
EI • ΔUit (/100)	.1098	.0985	.2603

	(0.453)	(0.226)	(0.099)
F-test value	6.43	6.66	7.06
R^2	0.11	0.11	0.12
Years	1997-2007	1997-2007	1997-2007
N	1538	1538	1538

Notes: The data was divided into highest years of schooling segments: elementary education (1-7 years), secondary education (8-12 years), post-secondary education (13-16 years) and graduate education (17+ years). Marginal values for the years of schooling are rounded up to the nearest integer. For example, an individual with 7.10 years of school is rounded up to have 8 years of education. Additional required definitions and notes are in table 2. P-values are in parenthesis.

Relative to age segmentation, segmentation by years of schooling gives more intuitive results. Overall, the direction and magnitude of the interaction terms were consistent at all years of schooling. The positive values of the interaction terms are consistent with the compensating differentials hypothesis. The magnitude and statistical significance are especially strong for the elementary and secondary school segments by the maximum EI measure. For the post-secondary and graduate sample group, only the qualifying EI measure is statistically significant at least at the ten percent level. Again, this is further evidence of composition bias, with elementary and secondary school education level individuals driving the results of the entire sample.

Table 9: Job changer wage cyclicality with income segmentation

	Maximum EI	Minimum EI	Qualifying EI
	<$10 Real Composite Hourly Wage		
Constant	2.0111	1.8759	1.8799
	(0.000)	(0.000)	(0.000)
ΔUit	-.0374	-.0058	-.0078
	(0.000)	(0.110)	(0.069)
EI • ΔUit (/100)	.0649	.0100	.0226
	(0.000)	(0.257)	(0.155)
F-test value	20.55	20.06	20.27
R^2	0.16	0.14	0.14
Years	1997-2007	1997-2007	1997-2007
N	8048	8048	8048
	$10 - $20 Real Composite Hourly Wage		
Constant	2.5137	2.4942	2.5012
	(0.000)	(0.000)	(0.000)
ΔUit	-.0155	.0010	-.0011
	(0.108)	(0.811)	(0.825)
EI • ΔUit (/100)	.0321	.0012	.0098
	(0.077)	(0.909)	(0.596)
F-test value	4.14	3.33	3.40
R^2	0.03	0.03	0.03
Years	1997-2007	1997-2007	1997-2007
N	5937	5937	5937
	$20 - $30 Real Composite Hourly Wage		
Constant	3.1317	3.1313	3.1358
	(0.000)	(0.000)	(0.000)
ΔUit	-.0060	-.0020	-.0036
	(0.470)	(0.598)	(0.409)
EI • ΔUit (/100)	.0107	.0040	.0124
	(0.495)	(0.659)	(0.450)
F-test value	1.85	1.78	1.88
R^2	0.02	0.02	0.02
Years	1997-2007	1997-2007	1997-2007
N	1764	1764	1764
	>$30 Real Composite Hourly Wage		
Constant	3.561	3.5280	3.5260
	(0.000)	(0.000)	(0.000)
ΔUit	-.0121	.0044	.0061
	(0.392)	(0.462)	(0.371)
EI • ΔUit (/100)	.0201	-.0142	-.0281

	(0.448)	(0.320)	(0.262)
F-test value	0.62	0.72	0.79
R^2	0.01	0.01	0.01
Years	1997-2007	1997-2007	1997-2007
N	492	492	492

Notes: The data was segmented into income levels based on real composite hourly wage in 1999 Canadian dollars. Additional required definitions and notes are in table 2. P-values are in parenthesis.

The regression results with income segmentation again focus on composition bias. The results are statistically significant at the five percent level only for job changers with composite hourly wages lower than ten dollars (1999$). Although the interaction terms of the other results were generally positive, they were statistically insignificant. Thus, lower income individuals are driving the results of the entire sample.

V. Conclusion

This paper sets out to resolve a mechanism to explain the difference in wage cyclicality between job changers and job stayers. Two major theories are tested using the Canadian Survey of Income and Labour Dynamics. First, the interindustry wage differentials hypothesis proposed by Okun (1973) is examined in which job changers earn higher wages during economic upturns due to higher rents and looser employment rationing by procyclical industries. Second, the compensating differentials hypothesis proposed by Barlevy (2001) is analyzed in which job changers face higher unemployment risk underlying their position in the high wage industry. Thus, job changers who move to higher wage jobs during economic upturns are merely compensated for future wage losses during economic downturns. The empirical results of this paper support the compensating differentials hypothesis. Furthermore, there is no evidence of the interindustry wage differential. However, the strength of this conclusion is dampened by evidence for composition bias. Certain sections of the sample possessing attributes of non-middle age, lower education, and lower income levels exclusively drive the results of the entire sample.

The conclusion offers some implications for policy makers and the academic study of economics as a whole. The composition bias evidenced in this paper reaffirms the dangers of treating labour markets as an aggregate of representative agents. Specification tests should always be performed for future studies as certain subgroups may be driving the results of the entire sample. This could lead to a flawed government policy. In this study, whereas the wages of individuals with higher incomes does not respond to EI levels of any measure when they change jobs, the wages of individuals with low income respond strongly to EI levels of every measure when they change jobs. Thus, government policies aimed at diminishing wage volatility through outside options such as EI would only affect lower income individuals. Lastly, the numbers of statistically insignificant results for the minimum and qualifying measures of EI levels imply that these values have been set too low. Thus, they may not offer real economic incentive such as bargaining power for job changers over the course of the business cycle.

VI. References

Abowd, John, and Ashenfelter, Orley. "Anticipated Unemployment, Temporary Layoffs, and Compensating Wage Differentials." In Studies in Labour Markets, edited by Sherwin Rosen. Chicago: University of Chicago Press for National Bureau of Economic Research, 1981.

Anderson, Patricia, and Meyer, Bruce. "Unemployment Insurance in the United States: Layoff Incentives and Cross Subsidies." *Journal of Labour Economics*, 11 (January 1993): S70-S95.

Barlevy, Gadi. "Why Are the Wages of Job Changers So Procyclical?" *Journal of Labour Economics*, 19 (October 2001): 837-78.

Beaudry, Paul, and DiNardo, John. "The Effect of Implicit Contracts on the Movement of Wages over the Business Cycle: Evidence from Micro Data." *Journal of Political Economy*, 99 (August 1991): 665-88.

Bils, Mark. "Real Wages over the Business Cycle: Evidence from Panel Data." *Journal of Political Economy*, 93 (August 1985): 666-89.

Murphy, Kevin and Topel, Robert. "Unemployment, Risk, and Earnings: Testing for Equalizing Wage Differences in the Labour Market." In Unemployment and the Structure of Labor Markets, edited by Kevin Lang and Jonathan Leonard. New York: Blackwell, 1987.

Okun, Arthur. "Upward Mobility in a High-Pressure Economy." *Brookings Papers on Economic Activity*, no. 1 (1973): 207-52.

Polsky, Daniel. "Changing Consequences of Job Separation in the United States." *Industrial and Labour Relations Review*, 52 (July 1999): 207-52.

Reder, Melvin. "The Theory of Occupational Wage Differentials." *American Economic Review*, 45 (December 1955): 833-52.

Shin, Donggyun. "Cyclicality of Real Wages among Young Men." *Economics Letters* 46 (October 1994): 137-42.

Solon, Gary; Barsky, Robert; and Parker, Jonathan. "Measuring the Cyclicality of Real Wages: How Important Is Composition Bias?" *Quarterly Journal of Economics*, 109 (February 1994): 1-26.

Stock, James; Wright, Jonathan; Yogo, Motohiro. "A Survey of Weak Instruments and Weak Identification in Generalized Method of Moments." *Journal of Business & Economic Statistics*, 20 (October 2002): 518-529.

Topel, Robert. "Equilibrium Earnings, Turnover, and Unemployment: New Evidence." *Journal of Labor Economics*, 2 (October 1984): 500-522.

Vroman, Wayne. "Worker Upgrading and the Business Cycle." *Brookings Papers on Economic Activity*, no. 1 (1977): 1229-50.

Vroman, Wayne. "Cyclical Earnings Changes of Low-Wage Workers." *Research in Labour Economics*, no. 2 (1978): 191-235.