UNION WAGE DISPERSION MAGNITUDE AND TREND BEFORE AND AFTER THE 2008 FINANCIAL CRISIS

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ABSTRACT

Over the past two decades the Canadian Labour market experienced a rise in the overall wage inequality and de-unionization. This paper provides empirical evidence on effects of unionization on the real hourly wage before and after the 2008 financial crisis in Canada. The Ordinary Least Square (OLS) model is used to estimate the impact of union status and time related (such as time trend, year, and post-2008) variables on the real hourly wage level. Moreover, Heckman's two-step approach is performed to correct for non-random sample selection. The findings show unionization affects the real hourly wage positively, while the time variables' coefficients indicate the 2008 financial crisis did not have a significant impact on the real hourly wage level in Canada.

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I. INTRODUCTION

Over the past several years, there has been a tremendous increase in income inequality within Canada and most advanced, emerging and developing countries (Heisz 2016). This reality has received considerable attention from the public, governments, and scholars alike. United States President Obama was quoted as saying, “controlling the widening global income inequality and understanding its impact in our society has become a defining challenge of our time”.[1] This new era of widening income inequality coincides with a decline in collective bargaining membership in most developed countries (Lemieux 1993). A number of studies point out that there is less wage dispersion among unionized workers than among comparable non-unionized workers in the United States and Canada (Card, Lemieux and Riddell 2004). The smaller wage dispersion in unions could be attributable to several factors. Unions have the power to reduce wage differentials related to the level of education and to the level of experience as well as to standardize wages among workers in the same industry and geographical location.

Unions’ stronger impact on the blue-collar sector compared to the white-collar sector increases the wage of the group at the lower end of the pay scale. Unions’ power to decrease wage dispersion among unionized workers is accompanied by an increase in the wage gap in comparison to non-unionized workers, which can in turn lead to a large variability in wages. These two opposing factors make it difficult to determine the overall impact of unions on the degree of inequality in the distribution of labour income in the economy. Some studies point out unionization’s equalizing effect on the degree of inequality in the distribution of income in the United States (Freeman 1980 and Hyclak 1979). DiNardo et al (1996) confirm that the decrease

[1] https://www.whitehouse.gov/the-press-office/2013/12/04/remarks-president-economic-mobility.

in unionization density in the U.S. has contributed to the increase in wage dispersion and a shrinking middle class. Comparative studies on unionization rate and level of income inequality for Canada and the United States show that the higher unionization rate in Canada accounts for 40% of the difference in income inequality between the two countries (Lemieux 1993). This suggests that unionization has a significant impact on income inequality.

Employment and Social Development Canada documents a decline in the unionization rate from 33.7% in 1997 to 31.5% in 2012.[2] It is a common belief that unions increase wages; however, unions also decrease the degree of income inequality among workers. The unions' impact on the degree of income inequality has been viewed differently pre/post 1980s. As explained by Blanchflower and Bryson (2004), there was a large focus on the monopoly aspect of trade unions before the 1980s. Johnson (1975) argued that the monopoly aspect enables trade unions to raise wages beyond the competitive level, creating and increasing the wage gap among unionized and non-unionized workers. In addition, the monopoly effect on wages and the rules and work restrictions accompanying unionization have negative impacts on output levels and productivity. On the other hand, Freeman (1980) presented results that were opposite of the previous belief by using micro data on unionized and non unionized workers. He concluded that the positive effects of unionization (reducing the variability in wages within unionized labour forces) were offset by the negative monopoly face presented previously.

It was not until the 1980s and with Freeman and Medoff's (1984) famous publication called "What Do Unions Do?" that economists started viewing unions as an arrangement to decrease the degree of income inequality in developed/developing societies. Freeman and Medoff

[2] http://www4.hrsdc.gc.ca/.3ndic.1t.4r@-eng.jsp?iid=17.

estimated the restraint of union monopoly power on output is as small as a 0.2%, and represents a 0.4% fall in GNP (Gross National Product).[3]

Economists, social scientists and politicians have become interested in the relationship between income inequality and financial crises. There have been a growing number of studies that suggest inequality weakens the growth, investment and increases financial instability over time (Bazillier and Héricourt 2015). Income inequality in Canada has increased over the past three decades. The after-tax income Gini coefficient has shown an increase since the mid-1990s (Heisz 2016). Reports published for the House of Commons and by the Conference Board of Canada show that the national income share for the highest and richest group has increased, while the share for the middle class and low-income group has decreased.

In this paper, I examine unions' effects on the real hourly wage of the Canadian workforce before, during, and after the 2008 financial crisis. For this purpose, I use the May and November Labour Force Survey data (LFS data) from 2006 to 2012. Moreover, I analyze the trend in wage disparity between unionized and non-unionized workers over the interval from 2006 to 2012. The results point out that unionization has a positive and statistically significant positive impact on real wages for the periods before and after the crisis.

This paper is organized as follows. Section II provides a brief summary of literature review based on Canadian, American and international papers in regards to the wage gap between workers presented by unions compared to non-unionized workers in the past few decades, as well as the effects of unions on overall labour income inequality. In section III, I present and

3 http://www.investopedia.com/terms/g/gnp.asp.

describe the data used and provide a summary of statistics. Section IV introduces the methodology used for this paper. Section V discusses the regression results. Lastly, section VI contains the conclusion, which summarizes the findings.

II. LITERATURE REVIEW

The overall relationship between unionization and income inequality has been of interest to many economists and social scientists over the past three decades. There have been many papers published in Canada and elsewhere over this period that focus on the union wage differential. These papers treat the effect on income inequality of unionization in Canada and internationally, which typically involves “between sector” and “within sector” wage effects associated with unionization cross-tabulated by gender, experience level, and education level using different data sets and different methodologies.

Employees are more likely to speak out or express their ideas, concerns and demands to their employers via collective action, which includes collective bargaining. One of a union’s main tasks is to protect their covered workers and ensure employers and governments hear their voice. Benefits and disadvantages of collective bargaining arrangements and its impacts on society and the economy have been discussed for decades. As noted by Bennett and Kaufman (2004), social scientists and economists considered two aspects for union arrangements: the monopoly face (i.e. the power to increase the wage above the free market, competitive level, and the concomitant decrease in the production level) versus the collective voice. Until the 1980s, economists and society mostly focused on the monopoly aspect, which predicted negative impacts on businesses and the economy through the channels of decreased profitability, higher labour costs, and reduced final output levels in unionized industries.

Checchi et al (2007) look at the connection between union membership and income inequality. They use several cross-national surveys from the International Social Survey Program (ISSP) for seven countries (West Germany, East Germany, Holland, Britain, Sweden, Norway, Italy) covering the period from 1985 to 2002. The ISSP conducts annual surveys on rotating topics such as religion, environment, income inequality, etc. In 1987, 1992 and 1999, the ISSP includes questions on social attitudes towards 1) government responsibility in transfer of income, 2) progressive taxation, and 3) level of acceptance for the current income inequality. The analysis shows that union membership is more concentrated in the middle earnings group. Increases in the income gap lead to a lower middle group and decreasing union membership over time. The second finding is in regards to workers' attitudes towards unions and collective bargaining. Unlike the United States and Canada, the selected countries have "open-shop"[4] arrangements (i.e. 'open-shop' is where all workers have wage and other protections in the selected countries). Therefore, the decline in union membership is due to union/collective agreement arrangements.

Lemieux and Riddell (2015) compared the income for top earners and the change in characteristics of the top 1% based on master files of the Canadian census over the period 1981-2006 and the National Household Survey (NHS) for 2011 (since the mandatory census long form survey was replaced by the voluntarily 2011 NHS). Census data have detailed demographic and socio-economic information, as well as labour-related information for the previous year.

4 "In an 'open-shop' framework, a single worker may observe the aggregate inequality measure as a proxy of the egalitarian attitude and effectiveness of unions, whereas in a 'closed-shop' context local inequality measures become crucial." Checchi et al (2007), pp. 13.

Furthermore, the educational attainment information is very detailed, which helps to examine the profile of the top 1% and which education group has contributed to the growth of the top group. Also, the census data is a large sample representing 20% of the Canadian population, which will produce a reasonable sample size from which to examine the education and occupation characteristics of the top 1%. This paper concludes that the income for top earners has grown faster than any other group over the past 30 years. Detailed education, industry, socio-economic and demographic information also shed light on how characteristics of top earners have changed over time in Canada. Similar to the case in the United States, executives and individuals in finance and business sectors are in the top 1% group. In Canada another main contributor to the growth in the 1% group is workers in the oil sector.

There have been many papers written about the impacts of unions in the economy citing Freeman and Medoff (1984). They discussed and measured the impacts of both aspects of unions and the decline in union density in their famous book "What Do Unions Do?" They asserted that productivity either stays neutral or increases in unionized environments. Unionized workers keep their jobs longer than non-unionized workers, and longer job tenures make unionized workers more productive. However, some economists argued that there is no increase in employer profitability because higher wages and benefits mean higher costs to employers, causing an increase in manager/employer resistance to collective bargaining formation. Sherk (2009) points out that unionized employers' profit is 15 to 20 percent less than comparable non-unionized companies. Freeman and Medoff conclude that the positive effects of unions on society and economy (lower resignation rate, longer job tenure, higher productivity, etc.) outweigh its negative impacts. They used several micro and panel data sets to analyze different aspects of unions. Using (mainly) post 1970s Current Population Survey (CPS) data, they concluded that unions have the impacts of reducing earnings inequality within the sector and throughout the labour market, as well as increasing other employee benefits (such as health insurance, pensions, vacations, etc.) within the sector.

Twenty years after Freeman and Medoff, Bennett and Kaufman (2004) agreed with the principal findings of "What Do Unions Do?" However, they expanded the initial findings by drawing on new empirical research done on unions, which has benefited from the availability of more complete data sets over the interlude. In addition, Bennett and Kaufman extended the scope of their analysis to cover union effects globally, since the original paper is based on American experience. Their findings are in line with those contained in Freeman and Medoff (1984).

Simpson (1985) measures and examines the impacts of unions on wages by using micro data from the 1974 Labour Canada Wage Survey. This paper examines the wage differential for different skill levels, industries and sectors (private versus public). It then compares the current findings to those contained in earlier papers, which show smaller union-non-union wage differentials than those reported in earlier studies. His results show that on average, a unionized worker's wage is 18.6% higher than a non-unionized worker (all other factors held constant). In addition, he finds the public-private sector wage differential is a result of higher union density in the public sector than in the private sector. However, selection bias may be present in determining the impacts of unions on wages, since the unionized labour force is not a random sample of the labour force. The Heckman procedure is used for adjusting selection bias.

Lemieux (1993) compares the effects of unionization on wage inequality in Canada and the United States. He uses the 1986-1987 longitudinal Labour Market Activity Survey (LMAS) file merged with the 1986 cross sectional file of the LMAS and the 1986 Current Population Survey for Canada and the United States, respectively. He shows that the effects of unions on wages are similar in Canada and the U.S. Unions reduce earning inequality for both genders within-sector. However, they decrease the overall wage gap for men while increasing the overall wage gap for women in Canada. The overall wage differential among men reduces to 14.5%, but the wage gap increases by 4.1% among women.

Riddell (1993) examines how the United States and Canada vary in terms of their degree of union organization, union densities, and the overall widening wage gap. This paper measures the union differentials and also investigates the factors behind the decreasing rates of unionization. During the 1920-1960 period, the two neighbours followed a similar pattern. However, after 1960 unionization started to decline in the United States, which led to a growing gap between the union density rates in the two countries. Therefore, this paper focuses on what caused this change between the two countries and measures the unionization differential between them. The data sets used are the Canadian Labour Force Survey (LFS) and the American Consumer Population Survey (CPS). Both surveys are similar in terms of coverage and methodology. This paper confirms that unionization declined in both countries between 1984-90, but much more rapidly in the United States. Union density in the United States declined in the period after World War II, but increased in Canada until the mid 1980s, regardless of individual worker characteristics; age, sex, education, etc. The proportion of Canadians represented by union/collective bargaining is twice as high as is the case in the United States (due mainly to the public sector).

Moreover, comparative analysis based on a supply and demand approach in the market for union representation shows that a Canadian worker who wishes to be represented by a union can eventually attain union status, but that is not true for the American worker. This is mainly due to the difference in the two countries' legal frameworks governing unions and collective bargaining. In addition, both countries' public sector employees are more likely to be unionized than are their private sector counterparts. In Canada, a higher fraction (about 30% more) of the working population is employed by the government entities than is the case in the United States, which is another contributor to the higher decline in unionization in the United States. The impact of unions on wages in the private sectors is similar in the two countries. Therefore, the difference in unionization differentials between the two neighbours can be attributed to differences in their legal regime and legislative environments concerning unionization.

DiNardo et al (1996) use hourly wage data from the U.S. Current Population Survey (CPS) and a semi-parametric procedure to analyze and investigate the institutional and labour market factors leading to changes in distribution of wages over 1979-1992 in the United States. Applying the weighted-kernel density method shows where these factors have the greatest impacts along the wage distribution. They show that fall in unionization and demand-supply shocks played an important role in wage inequality rise in the economy. In addition, unlike other previous studies, their analysis demonstrates that decreases in the real value of the minimum wage is another major contributor to the wage inequality increase within and between sectors in the United States over 1979-1988. Their findings show that wage inequality was less volatile during the 1973-1979 and 1988-1992 periods. The fall in the real value of the minimum wage in the 80s (approximately 25%) has a greater impact at the bottom end of the distribution (minimum wage earners, who are mostly women). They further disentangle the impacts of de-unionization, the decrease in real value of minimum wage, and demand and supply shocks on the wage distribution by applying the Oaxaca-Blinder method. Unlike the original Oaxaca method, which measures the gap in wage means of two groups, this paper decomposes changes in the wage distribution. They confirm that all three factors contributed to the expansion of wage inequality in the United States over the 1979-1992 period. They place a great importance on the de-unionization rise for contributing to the increasing wage inequality in the U.S. during the 1979-1992 period.

Card (1998) uses Current Population Survey (CPS) data on wages from May 1973 and 1974 plus the 12-month survey in 1993 to confirm that the level of unionization in the U.S. has fallen significantly over the past several decades, while the degree of wage inequality widened. Since the level of wage dispersion in a unionized/collective bargaining environment is smaller, there is a common belief that the fall in unionization/collective bargaining is one of the main contributors to the increase in income inequality. Thus, Card’s paper looks at the relationship between the decline in unionization and the rise in inequality for both genders in the United States over the period from 1973 to 1993. Card concludes that, for women, the increase in inequality cannot be explained by the fall in unionization, since the proportion of the female population represented by a union has been relatively stable since the early 1970s. However, male unionization declined over time, and this decline in unionization can explain about 15% to 20% of the observed increase in wage inequality among men. In addition, the increase in public sector unionization prevented a rise in the degree of wage inequality among women and men within this sector, while unionization declined significantly in the private sector.

Lemieux (1998) uses 1986-1987 longitudinal files of the Canadian Labour Market Activity Survey (LMAS) to analyze the impact of unions on wage levels and dispersion. The LMAS is preferred to standard panel data sets such as CPS because it includes information related to work history (such as job change) and it distinguishes between voluntary quits and layoffs. The unionized lower skilled workers receive a higher wage premium than their non- unionized counterparts. It is more likely for a lower skilled (observed skill) worker choosing and remaining with a unionized employer, while a higher skilled worker (observed skill) will have more incentive for leaving the union. Moreover, a non-unionized employer is less likely to hire a lower skilled worker who demands a higher than average wage premium. Therefore, union status depends on individuals' selection and employers, which leads to a non-random sample. Another weakness of the panel data is including observed skills only and not capturing the unobserved skills, which may impact wages considerably. Due to these shortcomings of panel data, Lemieux estimates a panel data estimator that corrects for selectivity bias and captures unobserved skills. This paper uses data on job changes to measure the unobserved productivity for unionized and non-unionized workers. The results show that unions on average increase wage levels and decrease wage inequality in Canada, and decrease returns for observed and unobserved skills.

Fang and Verma (2002) investigate the union versus non-union wage gap by using the 1999 Workplace Employee Survey (WES). This sample includes workers from both union (or collective bargaining) group and non-union group with data on earnings available (earnings are before tax, tips, overtime, bonuses and commissions). The earnings in the 1999 sample are compared with historical wage differentials between 1984 and 1998. According to their measures, the average unionized worker earned $20.36 in 1999, while the average non-unionized worker earned $17.82. This indicates a wage premium of approximately 14.3% before making any adjustment for individual, demographic, job, and workplace characteristics. However, after adjustment, the wage differential falls from 14.3% to 7.7%. The analysis concludes that the wage disparity is about 7.0%, which is lower than previous findings. This paper also shows a narrowing union-non-union wage gap, which could be due to decrease in the monopoly power of unions to set wages.

Using 30 years of micro data in the U.K., the U.S. and Canada, Card et al (2004) measure and compare the impacts of unions on wage inequality in these three countries. The data suggest that the unionization rate declined for all these three countries over time, with the greatest drop in the U.K., and the smallest drop in Canada. However, during the same period the degree of wage inequality increased for all three countries, with the smallest increase occurring in Canada and the largest increase occurring in the U.K. The analysis shows new evidence on the correlation between union density and the degree of earnings inequality in Canada, the U.K. and the U.S. Their analysis shows that unions have two offsetting factors on wage inequality as discussed in "What do Unions Do?". Unions increase the union-non union wage differential, which increases earnings inequality. On the other hand, unions reduce the wage gap among unionized workers, by increasing wages at the bottom of the income distribution. Wage dispersion reducing effects of unions among unionized employees dominate the union-non union wage inequality increasing effects. Similar to Freeman's (1980) findings, this paper concludes the "within sector effect" dominates the "between sector effect". Therefore, unionization decreases wage inequality.

The paper of Card et al (2004) draws on pooled May 1973 and 1974 Consumer Population Surveys (CPS) for the first wave of U.S. data. The CPS contains wage and union status information. For later years monthly earnings supplement files for 1984, 1993 and 2001 are used. The U.K. analysis is based on a combination of the 1983 General Household Survey (GHS) and the U.K. Labour Force Survey (UKLFS) (1993 & 2001). For measuring the impacts of unions in Canada, the authors use the Canadian Labour Force Survey (CLFS) for post 1997 (since they started asking wage and union questions in 1997), and for the period before 1997, the Survey of Unions Membership (1984) and Survey on Work Arrangements for 1991 and 1995. For salaried employees, the average hourly wage is estimated by dividing the weekly earnings by hours worked per week.

Green and Sand (2013) use Canadian census data and the Labour Force Survey (LFS) for the period 1971-2012 to examine whether the employment structure and wages have polarized, with lower employment and wage growth for middle-skilled workers compared to highly-skilled and low-skilled groups. Since most studies regarding this topic are based on American data, this paper uses U.S. census data as a benchmark for comparative purposes. The data sets used for this paper are Canadian census over the 1971-2006 period for available years (1971,1981, 1986, 1991,1996, 2001, 2006). The wage measure used for this study is weekly wages as in Card and Lemieux (2001). In order to look at the employment and wage trends beyond 2006, the LFS is used. As previously noted, U.S. data were used as benchmarks: the U.S. Census for 1970,1980, 1990 and 2000 and the American Community Survey for 2007. This paper finds that the Canadian labour market experienced polarization during the 1980s and 1990s, where employment grew for higher or lower skilled groups compared to mid-skilled occupations. However, this growth in employment started to stabilize for high-skilled workers post 2000. They found that the wage gap between the low-skilled and the middle-skilled group, as well as between the middle-skilled group and the high-skilled group started to rise.

Fortin and Lemieux (2014) use the Labour Force Survey (LFS) over the period 1997-2012 to take a more recent look at hourly wage inequality in Canada in all provinces. They use the LFS, since it contains questions on hourly wages, union status and firm size. Some of the analysis starts in 2009 due to changes in industry classification, which took place from 1998-99. Their results indicate that provincial wage dispersions are related to provincial minimum-wage changes, as well as faster growth in aggregate income for Alberta, Newfoundland and Saskatchewan due to the 1990s resource boom.

Acemoglu et al (2001) use previous studies on the relationship between de-unionization and wage inequality in the U.K. and the U.S. and propose a framework that links technical change to de-unionization. The fall in unionization and rise in wage inequality in the past twenty-five years bring attention to some explanations that may be primary contributors to the widening wage gap. The rise in wage inequality might be related to (i) labour market changes such as globalization and skill-biased technical change (ii) and institutional changes such as minimum wage and de-unionization. Data and results from earlier studies from Canada, the U.K. and the U.S. indicate that overall de-unionization lags wage inequality. This paper argues that when the skill-bias is larger than the benefit of unionization the skilled worker (who is in demand) will leave the union environment for higher pay, which leads to de-unionization. De-unionization is not the main cause of wage inequality but it strengthens the effects of skill-biased technical change by removing the wage compression imposed by unions. This paper considers de-unionization as a result of labour market changes (skill-biased technical change).

Fairris (2003) uses the 1984 and 1996 Mexican National Household Survey (INEGI) to measure and compare the impact of unionization on wage inequality in Mexico. The results show that trade unions had strong equalizing power on the dispersion of wages in 1984, but were only half as effective in reducing wage inequality by 1996. Over this period both the unionization density and the equalizing power were decreased considerably. This paper provides empirical evidence that the 1996 estimated formal sectors and overall wage inequality would have only decreased by 11% and 5.6%, respectively if the trade unions' ability on reducing the wage gap remained at the 1984 level. The results on the 1984 sample show a small and not statistically significant gender and geographical wage gap for unionized workers, when compared to non-unionized workers. The estimated coefficients for the 1996 sample indicate that the gender and geographical wage dispersion for union workers increased considerably and it is much closer to non-unionized workers.

Collective bargaining arrangement can indirectly impact the wage distribution via affecting economic growth. Chua et al (2016) examine the effects of a labour union in an open economy (two-country) R&D based economic growth model. In such model, innovation rates are determined by the domestic market size. The cross-country analysis is between the U.S. and the

U.K., since the U.S. is the largest economy in the world and the largest trading partner for the European Union. In addition, the U.K. data is chosen, since with respect to collective bargaining institutions the U.K. has the closest similarity to the U.S. among all European Union countries. They show that the increase in bargaining power of unions can affect domestic wages both negatively and positively depending on where the focus of bargaining union lies. Increase in bargaining power of a wage-oriented union affects domestic employment negatively, which puts a downward pressure on domestic innovation and economic growth. Moreover, it causes the level of innovation in the foreign country to go up which has negative effect on domestic wages compared to foreign wages. However, increase in bargaining power of an employment union will positively affect employment and increase wages for the domestic country compared to the foreign country. They use OECD data from 1980 to 2007 for trends of wage share in GDP growth. They consider U.S. Bureau of Labor Statistics and U.K. Office for National Statistics from 1980 to 2007 for trends of the unemployment rate between the two countries. Post 2007 data is not considered due to the financial crisis.

Most previous studies focus on the distribution of the rising wage inequality in the past several decades with not much attention paid to the trend in real wage growth. Machin (2016) links the increasing wage inequality and the slowing real wage growth. He uses OECD data between 1980 to 2013 to analyze the increasing wage inequality in U.K. and U.S. Data for analyzing the trend in the real wage growth in the U.S. comes from the Current Population Survey (CPS) from 1980 to 2013. The comparison between the wage inequality trend and slowing real wage growth indicates that the real wage stagnation lagged increasing wage inequality. The U.K. real wage continued to grow at 2% until the mid-2000s; however, the growth started to slow down and turned negative after the 2008 financial crisis. This paper provides empirical evidence that there is a connection between the rise in wage inequality and the real wage stagnation in the two countries.

From 1973 to 2007 the American private sector union membership decreased from 34% to 8% and 16% to 6% for men and women, respectively. Most studies focus on the effects of the unionization rate on the union wage dispersion or the overall wage inequality. Rosenfeld and Western (2011) use the Current Population Survey (CPS) to analyze the relationship between the fall in unionization and its impacts on the private sector wage dispersion among unionized industries. The findings indicate that unionization's wage equalizing power continues to hold in the private sector for industries that are heavily unionized. They use the variance decomposition method to measure the growth in hourly wage among private sector full time employees. The results show that the fall in unionization contributes 20% of the increase in wage dispersion for private sector workers. This paper uses annual May files of CPS from 1973 to 1981, and the annual merged Outgoing Rotation Group files of CPS from 1983 to 2007. The 1982 data is excluded due to its lack of union status question. Also, the 1994 and three-quarter of 1995 data are not included because the wage allocation flag was missing for private sector employees.

In recent decades much attention has been paid to the effects of unionization on wage distribution, while there is a small body of research that focuses on the impacts of unionization on income distribution. Herzer (2014) uses the Estimated Household Income Inequality (EHII) database developed by the University of Texas Inequality Progress (UTIP) to evaluate the relationship between unionization and income inequality across twenty countries as a whole and each country individually. Employing heterogeneous panel cointegration technique determines whether there exists a relationship between the level of unionization and the level of income inequality, as well as the direction of long-term causality across countries. This type of analysis shows the average impact of trade unions on income inequality and allows to examine the country differences. Unions can impact the distribution of income inequality through (i) affecting the wage distribution (ii) affecting employment (iii) and affecting the wage share. Previous papers show the effect of unionization on income distribution can be positive, negative or unknown. The analysis shows that (i) unions on average widen the income inequality level in the long run, (ii) there are large differences on impacts of unions on income inequality, in 65% of the countries decrease in unionization level increases income inequality, while in the 35% remaining cases fall in unionization rate positively impacts the income inequality, (iii) and long-run causality between the level of unionization and the level of income inequality runs both directions, which means in general the fall in unionization level increases income inequality and rises in income inequality leads to decrease in the level of unionization.

Unionization has the power to reduce the degree of wage dispersion and income inequality. As previously discussed wage and income inequality have been rising within Canada and globally over the past few decades, while the unionization rate has been declining. After the most recent financial crisis, economists and politicians have become interested in the relationship between income inequality and the 2008 financial crisis. Some believe that the high degree of income inequality is a possible contributing factor to the credit/financial crisis. When income inequality rises, individuals tend to raise their spending and finance it by debt in order to maintain the same level of consumption and living standards, which is a significant contributor to credit bubbles.

Based on existing literature, Bazillier and Hericourt (2015) present and overview of the two-way relationship between income inequality and the financial crisis. There is strong evidence that supports the role of income inequality in the credit bubble and financial crisis. Growth in the degree of income inequality increases the demand for credit and already easily accessible credit will increase indebtedness.

In this paper, I focus on effects of unionization on wage dispersion over the period before, during and after the 2008 financial crisis.

III. DATA

In order to examine whether there is any relationship between the 2008 financial crisis and the union wage differential, I start by drawing data from the May and November Labour Force Survey (LFS) over the period 2006 to 2012. The LFS is a monthly survey with cross-sectional design, which includes detailed information required for the analysis, such as union status, sex, marital status, employment sector, education attainment, age category, hourly wage, and province of employment. The LFS surveys the respondents for six consecutive months; therefore, in order to ensure independent observations, for every year only two months of LFS data were used (May and November). The two monthly surveys for 2006, 2007, 2008 and May 2009 were combined to construct a pre-2008 sample, and the post-2008 period sample combines November 2009 with May and November for each of 2010, 2011 and 2012. The population of interest for the LFS is the Canadian population age 15 and older in all provinces and territories, excluding full-time members of the Canadian Armed Forces, institutionalized people, and persons living on reserves or other aboriginal settlements. The overall number of observations is

1,461,280; with 616,769 observations belonging to the pre-2008 period (including May 2009) and 844,511 observations over the post-2008 period November 2009-2012.

The sample data for analysis employed to measure the impacts of collective bargaining on wage differentials for the pre-2008 period and the post-2008 period are restricted to the employed population aged 20 to 65 years. This includes persons in the indicated age group who are employed during the survey week, excluding observations for which the union information is missing and self-employed individuals, since there is no available hourly wage information for this group. I chose age 20 as the lower limit, since the number of observations for unionized members younger than that age category is small. I set the upper limit at 65 years, which is the normal retirement age in Canada.

Although the LFS data include information for most explanatory variables, the experience/skill information is not included. Potential experience as calculated according to the Mincer way (experience is age minus the number of years in school minus 6 as the school-start age) is not possible to determine as the precise age and the total number of years of schooling are not provided. Information on both of these variables is grouped into educational categories. For this paper, I use the method developed in Card et al (2004), which used the educational attainment group and the age group as proxy variables for educational attainment and age.

The oldest age group is comprised of those aged 50-64 (i.e. not a 10 year category in order to increase the number of observations for that group), since the majority of union workers benefit from more generous pension plans arrangements and is often associated with early retirement.

The number of observations for the age group 50 to 65 decreases significantly when compared to non-unionized workers in the same age group. Working beyond retirement has increased after the financial crisis among non-unionized workers.

Table 1 provides the before (and during) and after crisis weighted means and standard deviations for all of the explanatory variables. The union mean indicates that proportion of the union membership during the post financial crisis period has decreased by 0.2 percentage points to 32.8% compared to 33.0% for the period after the crisis. For both periods the unionized proportion that I calculated is slightly higher than the one reported by Statistics Canada in 2012: rate of 31.5%.[5] Note that in my sample, the individuals covered under collective agreements are treated as unionized. The proportion of public sector workers increased from 24.9% (before the financial crisis) to 25.7% (after May 2009), although the unionization rate decreased.

The female shares for before and after the 2008 financial crisis are 49.5% and 49.8%, respectively. As far as gender is concerned, this sample seems to be representative, since the proportions of males and females are about 50%. The three statistics representing marital status show minor fluctuations in the means for the two periods under analysis. The percentage of either married or common-law individuals remained at about 65% over the period from 2008 to 2012. The single (never married) category represents 26.3% of the Canadian population before the crisis and 26.6% after 2008. The divorced/widowed/separated group represents about 9.0% over the pre-May 2009 period. This group shrank in size by 0.2 percentage points after May 2009.

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5 http://www.labour.gc.ca/eng/resources/info/publications/union_coverage/union_coverage.shtml.

The categorical education variable divides the sample into three levels of educational attainment, with a majority of the population either having a college degree or some post-secondary education. Table 1 shows that about 37.4% of Canadians had at least some post-secondary education, which slightly increased post crisis by 0.5 percentage points. The statistics for university graduates also show an increase for the period after the financial crisis. The university graduate shares for before and after 2008 are 24.9% and 27.2%, respectively. I pool together high school dropouts and high school graduates following the procedure in Lemieux (1998). The population of high school (or less) graduates decreased slightly during the post 2008 period. The statistics presented in Table 1 are consistent with the decrease in the young working population and an increase in the mature working population.[6]

The age variable is divided into four sub-groups; 20-29 years, 30-39 years, 40-49 years and 50-64 years. For the period before the 2008 crisis, the age group of 40-49 represents approximately 27.3% of the working population, which is a higher share than any of the other three groups. After the crisis occurred, however, it is the 50-64 age group that comprises the highest share. The sample share for the group 50-64 has increased from 24.6% to 27.0%. Except the oldest age group, the sample size for the other three age groups decreases over time (pre-2008 versus post-2008). The changes in the share for the period before and after the crisis for the four age groups are consistent with the aging population in Canada.

[6] https://www12.statcan.gc.ca/census-recensement/2011/as-sa/98-310-x/98-310-x2011003_1-eng.cfm.

The regional category shows the share of each province in total, with approximately 39% of the population in Canada residing in Ontario and 23.3% in Quebec. The explanatory variable of marital status remained approximately unchanged from 2006 to 2012.

IV. EMPIRICAL METHODOLOGY

I use weighted linear regression to estimate an equation whose dependent variable is the logarithm of the real hourly wage and to compare some of the wage determinants over the period from 2006 to 2012. The first regression is the baseline model, which measures the overall impact of unionization on the logarithm of real hourly wage over the period 2006 to 2012. It takes the form:

$$\ln(\text{real-}W_i) = \beta_0 + \beta_1 \text{ union}_I + \varepsilon_i \quad \textbf{(1)}$$

where ln (real-W_i) is the natural logarithm of real hourly wage for individual i. Union i is the variable of interest representing union status for individual i. For the second regression (illustrated below), I added a post financial crisis dummy variable into the baseline regression. Adding a post financial crisis variable determines whether the impact of unionization on the wage evolved over this period.

$$\ln(\text{real-}W_i) = \beta_0 + \beta_1 \text{ union}_i + \beta_2 \text{ post-2008}_i + \beta_3 \text{ union}_i * \text{post-2008}_i + \varepsilon_i \quad \textbf{(2)}$$

For the third specification, I added demographic and geographical variables to equation (2) to measure the impacts of each of those explanatory variables on the natural logarithm of the real hourly wage. The control variables are all binary variables: union status, post-2008, public-sector, female, marital status (with three binary variables), education level (three binary variables), age group (four binary variables), and location (labeled P dummies for nine provinces j=1…9) to measure the impacts of each of those explanatory variables on natural logarithms of the real hourly wage.

$$\ln(\text{real-}W_i) = \beta_0 + \beta_1 \text{ union}_i + \beta_2 \text{ post-2008}_i + \beta_3 \text{ public-sector}_i + \beta_4 \text{ female}_i + \beta_5 \text{ common-law}_i + \beta_6 \text{ divorced /separated/widowed}_i + \beta_7 \text{ college}_i + \beta_8 \text{ university}_i + \beta_9 \text{ age\_20}_i + \beta_{10} \text{ age\_40}_i + \beta_{11} \text{ age\_50\_65}_i + \sum\nolimits^{9}_{j=1} P_{ij}\, \alpha_j + \varepsilon_i \quad \textbf{(3)}$$

Further, I expanded the third regression to include interaction terms between the union status variable and other exogenous variables to examine the heterogeneity in the effect of union status across the control variables.

$$\ln(\text{real-}W_i) = \beta_0 + \beta_1 \text{ union}_i + \beta_2 \text{ post-2008}_i + \beta_3 \text{ public-sector}_i + \beta_4 \text{ female}_i + \beta_5 \text{ common-law}_i + \beta_6 \text{ divorced /separated/widowed}_i + \beta_7 \text{ college}_i + \beta_8 \text{ university}_i + \beta_9 \text{ age\_20}_i + \beta_{10} \text{ age\_40}_i + \beta_{11} \text{ age\_50\_65}_i + \sum\nolimits^{9}_{j=1} P_{ij}\, \alpha_j + \beta_{12} \text{ union*post-2008}_i + \beta_{13} \text{ union*public-sector}_i + \beta_{14} \text{ union*female}_i + \beta_{15} \text{ union*common-law}_i + \beta_{16} \text{ union*divorced /separated/widowed}_i + \beta_{17} \text{ union* college}_i + \beta_{18} \text{ union* university}_i + \beta_{19} \text{ union* age\_20}_i + \beta_{20} \text{ union* age\_40}_i + \beta_{21} \text{ union* age\_50\_65}_i + P_i\, \alpha_i\text{*union} + \varepsilon_i \quad \textbf{(4)}$$

Working for a unionized employer can be a choice that an employee makes based on personal circumstances and preferences. This implies that the sample used for our regression is not random, which causes selection bias. I use the Heckman two-step method to correct for selectivity bias in this paper. The first step calculates the probability of union status. The second step consists of the wage equation that includes the inverse mills ratio term.

$$\ln(\text{real-}W_i) = \boldsymbol{\beta}\, \mathbf{X_i} + \varepsilon_{reg\_i} \qquad \textbf{(5)}$$

Where $\mathbf{X_i}$ is the vector of previously described explanatory variables. The individual's probability function for selecting union status is as follows:

$$P_i = \boldsymbol{\mu}\, \mathbf{Z_i} + \varepsilon_{i\_selection.} \qquad \textbf{(6)}$$

$\mathbf{Z_i}$ in equation (6) is a vector of all variables that impact the union status selection. An individual is unionized if $\boldsymbol{\mu}\, \mathbf{Z_i} + \varepsilon_{i\_selection} > 0$. Matrices **X (**in equation 6) and **Z** (in equation 5) may have variables in common. However, there should be at least one variable in vector **Z** that has an impact on choosing union status but does not impact the hourly wage and is excluded from vector **X**. The control variables, which are included in the discrete choice equation are binary variables: female, marital status, education level, and age group. In addition the matrix **X** includes an instrumental variable, which is the spouse's employment status, which I assert affects the individual decision on selecting union status but has no impact on the hourly wage.

When a worker is the main earner in the family (his/her spouse is unemployed and searching for work), the probability of that individual's preference to a more secure job is higher. The correlation between spouse's employment status and the hourly wage is close to zero for the pre 2008 and the post 2008 samples. However, there is positive correlation between the employment status of the spouse and holding union jobs.

The sample used for Heckman two-stop procedure excludes single employees, so it is smaller than the sample used for OLS estimation. By taking conditional expectation on equation (6) with respect to the vector $\mathbf{X_i}$ and P_i=1, we can determine that the value for ε_{reg_i} (error term in equation 5) is not zero.

The second step in the Heckman method corrects the wage equations for selection bias. The top section of Table 3 presents the results after correcting for the selection bias for each regression. There is a minimal change in adjusted results (in Table 6) when compared to the estimated coefficients of the unadjusted regression coefficients. As will be explained in section V, the inverse mills ratio term is insignificant, indicating that the Ordinary Least Squared (OLS) estimates obtained from the unadjusted regressions do not need to be adjusted.

V. RESULTS

The results of the regression equations for the period from May 2005 to November 2012 are displayed in Table 2, which reports the unadjusted effects of all explanatory variables on the logarithm of the real hourly wage. As shown in Table 1, the unionization density for both periods remains about 33.0%, which is slightly higher than the value of 31.5% reported by Employment and Social Development Canada in 2012. The post-2008 real average hourly wage is slightly higher than the pre-2008 period. The real average hourly wage increased from \$21.08 to \$21.90 after the financial crisis.

From the unadjusted regression results presented in Table 2, it is apparent that unionization has a significant impact on the real hourly wage under all scenarios (columns i-iii). Holding all other variables constant, a unionized worker enjoys a 14.0%[7] higher real hourly wage on average when

[7] exp(0.131)-1=14.0%.

compared to the reference group (the base group is a non-unionized, private sector single male, with no post secondary education, aged 30-39, and working in British Columbia during the pre-2008 period). The union status variable is statistically significant at the 1% level. In table 3, the post-2008 time variable is replaced by six dummy variables representing every year (using 2007 as the reference year). As shown in table 3, being a unionized worker positively impacts the real hourly wage. Holding all other variables constant, being unionized on average raises the real hourly wage by 14.5% and the union status variable is significant at the 1% level.

Similarly, linear trend regression results presented in Table 4 confirm the statistical significance of union status discussed above. Table 5 contains the estimates for the impacts of union status on real hourly wage pre-2008 and post-2008 financial crisis. The union's status power over real hourly wages increased from 12.7% to 14.3% during the later period.

Besides union status, variables such as the time trend, public sector status, gender, marital status, age, education and location of employment have impacts on wage. The following paragraphs explain the impacts of these explanatory variables on the real hourly wage.

<u>The impact of the Sector of Employment</u>

As shown in table 2(iii), being a public sector worker positively impacts the real hourly wage. Holding all other variables constant, being employed in the public sector on average raises the real hourly wage by 22.1% (table 2 column iii, table 3 and table 4). Table 5 compares the impact of public sector on the real hourly wage over two periods: pre-2008 and post-2008 financial crisis. The public sector wage-premium increased from 18.9% to 21.9% during the post-2008 period. The estimated coefficient of the public sector variable is statistically significant at the 1% level under all of the specifications.

The impact of the 2008 financial crisis and the Time-trend

As shown in table 2(ii) and 2(iii), the 2008 financial crisis has positive impacts on the real hourly wage. Holding all other variables constant, during the post financial crisis period, on average the real hourly wage rose by 2.74% (table 2, column iii). Moreover, table 3 presents the impacts of year dummy variables on the real hourly wage over the period 2006-2012. The year of 2006 is the only year that displays negative impact on the real hourly wage relative to the reference category of 2007. However, the overall impact of the time trend variable on the hourly wage is minimal in magnitude and is less than 1% (as shown in table 4). The time trend and post-2008 explanatory variables are statistically significant at the 1% level.

The impact of Gender

Column (iii) in table 2 shows that being a female worker negatively impacts the hourly wage by 20.2%. The same result holds for regression results presented in tables 3 and 4. Moreover, the female coefficient results in table 5 indicate that the negative impact of female status on real hourly wages decreased from 20.9% to 19.5% during the post-2008 period. The findings (under all specifications) imply that the hypothesis of no gender disparities can be rejected at the 1% level of significance.

The impact of Marital Status

Results presented in table 2 column (iii), table 3 and table 4 show that living in conjugal relationship (presently or in the past) impacts the hourly wage positively. Holding all the other variables constant, a married (or in a common law relationship) worker earns (on average) 12.1% more than a single employee. Similarly, having separated/widowed status increases the real hourly wage by 7.26% compared to the wages of a single worker. Furthermore, table 5 shows that the impact of marital status on the real hourly wage has slightly increased during the post financial crisis period for a unionized worker. Marital status variables are statistically significant at the 1.0% level.

The impact of Educational Attainment

Results presented in table 2 column (iii), table 3 and table 4 confirm that a higher level of education impacts the hourly wage positively. For a unionized female working after the financial crisis, raising the level of educational attainment from no post secondary education (high school or less) to a college (or some post secondary) degree increases the real hourly wage (on average) by 17.5%. Similarly attainment of a university degree increases the real hourly wage by 44.7% when compared to a worker with no post secondary education. However, the results presented in table 5 show that the returns to higher education on the hourly wage have slightly decreased after the financial crisis for a unionized worker. The estimated coefficients of the variables indicate that the level of education is statistically significant at the 1.0% level.

The impact of Age

The results contained in table 2 column (iii), table 3, and table 4 demonstrate that a young unionized worker (aged 20-29 years) earns on average 18.0% less than a non-unionized male worker between 30 to 39 years old. It is important to note that older age usually is associated with more experience, and the wage is usually expected to increase as age increases. As per table 2 column (iii), a unionized worker aged 40 to 49 years earns on average 5.0% more when compared to the base group (my reference category is not unionized and aged 30 to 39 years). For a unionized worker between 50-64, the wage premium increase is 3.3% when compared to non-unionized 30-39 year-old workers. This increase is less than the increase for 40-49 years old unionized workers. This could be due to most unionized workers benefiting from more generous pension arrangements and retiring (dropping out of the sample). Based on the above analysis, I conclude that the age variable plays an important economic role. Categorical age variables are statistically significant at the 1% level.

The comparison between the pre and post financial crisis periods presented in table 5 confirms the same pattern for impacts of age on the hourly wage. However, the results indicate that the age related wage premium increased slightly for older workers after the financial crisis. The comparison between the two periods shows no significant change for the youngest group.

The impact of Province of Residence

Lastly, the province of residence is another explanatory variable that has an impact on the average hourly wage. The findings contained in table 2, column (iii), table 3, and table 4 display large variability across provinces in wage differentials among unionized workers compared to non-unionized workers relative to the union wage premium in BC. Except for the case of Alberta, all provinces confirm that the province of employment impacts the real wage negatively relative to BC. Performing hypothesis testing indicates that the estimated coefficients of the province of employment are significant at the 1% level (with the exception of Ontario being significant at the 5% level).

Results presented in Table 5 indicate that the negative provincial impacts on the hourly wage (compared to the reference group) decreased during the post 2008 period, while in Alberta the wage differential became more positive. The most significant change took place for the province of Saskatchewan, where the provincial wage differential increased from negative 5.2% to positive 2.1%.

As evident in other studies, variables such as actual age (not being able to use age squared), experience, the participant's view on collective bargaining, and years of education are other determinants that are important in explaining the union wage premium. However, I could not use these variables in my analysis, since the LFS does not provide such information. Therefore, I have a problem of endogeneity (stemming from omitted variables) since some significant determinants are not included as independent variables in the regression equations for determining the dependent variable.

Application Of The Heckman Sample Selection Bias Correction Procedure

As previously discussed in section III, the sample of unionized workers used for my regression is not random, and this causes selection bias as far as estimating the union wage premium is concerned. I use the Heckman's two-step procedure on non-single sample for determining and correcting for selection bias.[8] Table 6 presents the estimated probability of a participant being unionized and estimated effects of explanatory variables on the logarithm of the real hourly wage after correcting for sample selection bias. As shown in table 6, the coefficients of the inverse mills ratio for the pre-2008 and post-2008 periods are 0.04 and -0.04, respectively. Moreover, both coefficients are insignificant, so the selection bias is minimal and as a result Ordinary Least Square (OLS) estimation can be used to estimate the hourly wage equation. A positive value for the inverse mills ratio coefficient during the pre-2008 period means that having an unemployed spouse has a positive effect on selection into unionized work, and a positive rho value positively impacts the hourly wage. On the contrary, the negative inverse mills ratio coefficient during the post-2008 period indicates that having an unemployed spouse impacts the selection into unionized work negatively.

[8] The instrumental variable used is the spouse's status of employment, which could affect the individual's choice on selection of union job but not the hourly wage earned. Therefore, single employees are excluded in Heckman's two-way procedure.

VI. CONCLUSION

This paper analyzes the effects of various factors on the logarithm of the real hourly wage. As has been discussed in previous literature one expects that unionization would have a positive and significant effect on the hourly wage during pre and post 2008 financial periods.

This paper uses the May and November waves of the Canadian Labour Force Survey (LFS) over the period 2006 to 2012 to estimate an earnings function that includes a union variable with an emphasis on the impacts of unions on wages. In addition, I explore the pattern of unionization's impact on the hourly wage for the period before and after the 2008 financial crisis.

Variables such as gender, age, level of education, marital status, and province of employment are some of the variables included in the LFS data that impact the hourly wage. However, for the purpose of this study, I mostly focus on the impact of unionization on the logarithm of the hourly wage. I also analyze the interactions of other explanatory variables with union status. Analyzing the union variable's coefficients under all specifications shows that unionization has a positive and significant impact on the logarithm of real hourly wage during both periods. Overall, a unionized employee (on average) earns 14.0% higher in wages when compared to a non-unionized employee. The small value of the coefficients for the time trend variables and the year dummy variables indicate that the fluctuations in the impact of unionization on earnings remained minimal in Canada over the period 2006 to 2012.

There are some demographic characteristics that have effects on selection into unionized work. Therefore, I estimated the hourly wages by performing Heckman two-step procedure to test for sample selectivity. Performing hypothesis testing on the coefficient of the inverse mills ratio indicates that there is no selection bias. Therefore, Ordinary Least Square (OLS) estimations calculated in unadjusted regressions are valid, and there is no need to correct for selection bias.

In conclusion, my findings on the importance of unionization on earnings agrees with results published in previous studies. The minimal fluctuation in the estimated coefficients of the time variables on the logarithm of the hourly wage over the period 2006 to 2012 suggests that the financial crisis did not have a significant impact on hourly wages in Canada.

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Table 1

Summary Statistics: Mean and Standard Deviation

	Pre-2008 Mean		**Post-2008 Mean**	
1) Union Status				
Unionized	0.332	(0.471)	0.330	(0.470)
2) Sector				
Public	0.249	(0.432)	0.257	(0.437)
3) Gender				
Female	0.495	(0.500)	0.498	(0.500)
4) Marital Status				
Married or Common Law	0.647	(0.478)	0.647	(0.478)
Separated, Widowed, Divorced	0.090	(0.286)	0.088	(0.283)
Single (never married)	0.263	(0.440)	0.265	(0.442)
5) Educational attainment				
High School or Less	0.378	(0.485)	0.350	(0.477)
College or some Post Secondary	0.373	(0.484)	0.378	(0.485)
University Graduate	0.249	(0.432)	0.272	(0.445)
6) Age Group (years)				
20 to 29	0.241	(0.428)	0.239	(0.426)
30 to 39	0.239	(0.426)	0.236	(0.425)
40 to 49	0.273	(0.445)	0.253	(0.435)
55 to 64	0.247	(0.431)	0.272	(0.445)
7) Province				
Newfoundland	0.014	(0.117)	0.014	(0.118)
Prince Edward Island	0.004	(0.064)	0.004	(0.064)
Nova Scotia	0.028	(0.164)	0.027	(0.162)
New Brunswick	0.022	(0.148)	0.021	(0.145)
Quebec	0.232	(0.423)	0.233	(0.423)
Ontario	0.394	(0.489)	0.390	(0.488)
Manitoba	0.036	(0.186)	0.036	(0.187)
Saskatchewan	0.028	(0.169)	0.029	(0.168)
Alberta	0.115	(0.319)	0.118	(0.323)
British Columbia	0.126	(0.332)	0.127	(0.333)
8) Income (constant 2006 dollars)				
Hourly Real-Wage	21.08		21.90	
Number of Observations	343,102		346,027	

Notes: Sample data is restricted to employed workers reporting a union sector indicator (not self employed) between 20 and 65 years of age. The result is based on sample data (May & November between 2006 to May 2009 for Pre-2008 data and November 2009 to 2012 for Post-2008 LFS). All means are weighted.

Table 2

Regression Model: logarithms of real hourly wage including the union status variable, pre/post 2008 time variable, and categorical variables.

Explanatory Variables	**(i)**		**(ii)**		**(iii)**	
Union	0.233	(0.001)***	0.231	(0.002)***	0.131	(0.005)***
Post 2008			0.039	(0.002)***	0.027	(0.002)***
Public					0.203	(0.003)***
Female					-0.226	(0.002)***
Married or Common Law					0.114	(0.002)***
Separated/Widowed/Divorced					0.071	(0.004)***
University Graduate					0.370	(0.003)***
College/some post secondary					0.161	(0.002)***
20 to 29 years					-0.199	(0.003)***
40 to 49 years					0.053	(0.003)***
50 to 64 years					0.031	(0.003)***
Newfoundland					-0.215	(0.004)***
Prince Edward Island					-0.254	(0.004)***
Nova Scotia					-0.184	(0.004)***
New Brunswick					-0.213	(0.003)***
Quebec					-0.086	(0.003)***
Ontario					-0.005	(0.003)***
Manitoba					-0.098	(0.003)***
Saskatchewan					-0.015	(0.003)***
Alberta					0.134	(0.003)***
Interactions of Union with:						
Post 2008			0.003	(0.003)***	0.008	(0.003)***
Public					-0.021	(0.004)***
Female					0.063	(0.003)***
Married or Common Law					-0.041	(0.003)***
Separated/Widowed/Divorced					-0.018	(0.005)***
University Graduate					-0.067	(0.004)***
College/some post secondary					-0.027	(0.003)***
20 to 29 years					0.046	(0.004)***
40 to 49 years					-0.018	(0.004)***
50 to 64 years					0.007	(0.004)***
Newfoundland					0.093	(0.006)***
Prince Edward Island					0.095	(0.006)***
Nova Scotia					0.047	(0.005)***
New Brunswick					0.076	(0.005)***
Quebec					-0.003	(0.004)***
Ontario					0.014	(0.004)***
Manitoba					-0.023	(0.005)***
Saskatchewan					-0.029	(0.005)***
Alberta					-0.079	(0.005)***
Intercept	2.873		2.852		2.777	
R-squared	0.051		0.053		0.299	
Observations	689,129		689,129		689,129	

*Notes: Logarithm of Real hourly Wage is the dependent variable. Survey weights are utilized. Standard Errors are presented in parenthesis. The explanatory variables are union, post financial crisis indicator, public sector status, categorical dummy variables (sex, marital status, education and age) plus nine provincial dummies, in addition to interaction all variables with union status. The base group is non-unionized pre 2008 private sector single male, with no post secondary education, age 30-39 working in British Columbia. Significance level: Significance level: * p<0.10 ** p<0.05 *** p<0.01 in a two-tailed test.*

Table 3

Estimates of a regression model: logarithm of the real hourly wage including annual year variables.

Explanatory Variables		
Union	0.135	(0.006)***
Year-2006	-0.013	(0.003)***
Year-2008	0.020	(0.003)***
Year-2009	0.041	(0.003)***
Year-2010	0.036	(0.003)***
Year-2011	0.028	(0.003)***
Year-2012	0.034	(0.003)***
Public	0.202	(0.003)***
Female	-0.226	(0.002)***
Married or Common Law	0.114	(0.002)***
Separated/Widowed/Divorced	0.071	(0.004)***
University Graduate	0.370	(0.003)***
College/some post secondary	0.161	(0.002)***
20 to 29 years	-0.199	(0.003)***
40 to 49 years	0.053	(0.003)***
50 to 64 years	0.031	(0.003)***
Newfoundland	-0.214	(0.004)***
Prince Edward Island	-0.253	(0.004)***
Nova Scotia	-0.184	(0.004)***
New Brunswick	-0.213	(0.003)***
Quebec	-0.086	(0.003)***
Ontario	-0.005	(0.003)***
Manitoba	-0.098	(0.003)***
Saskatchewan	-0.015	(0.003)***
Alberta	0.134	(0.003)***
Interactions of Union with:		
Year-2006	-0.001	(0.005)***
Year-2008	-0.006	(0.005)***
Year-2009	-0.008	(0.005)***
Year-2010	0.010	(0.005)***
Year-2011	0.004	(0.005)***
Year-2012	0.001	(0.005)***
Public	-0.021	(0.004)***
Female	0.062	(0.003)***
Married or Common Law	-0.041	(0.003)***
Separated/Widowed/Divorced	-0.018	(0.005)***
University Graduate	-0.067	(0.004)***
College/some post secondary	-0.027	(0.003)***
20 to 29 years	0.046	(0.004)***
40 to 49 years	-0.018	(0.004)***
50 to 64 years	0.008	(0.004)***
Newfoundland	0.093	(0.006)***
Prince Edward Island	0.095	(0.006)***
Nova Scotia	0.047	(0.005)***
New Brunswick	0.076	(0.005)***
Quebec	-0.003	(0.004)***
Ontario	0.014	(0.004)***
Manitoba	-0.023	(0.005)***
Saskatchewan	-0.029	(0.005)***
Alberta	-0.079	(0.005)***
Intercept	2.777	
R-squared	0.299	
Observations	689,129	

*Notes: Logarithm of Hourly Wage is the dependent variable. Survey weights are utilized. Standard Errors are presented in parenthesis. The explanatory variables are union status, time trend, public sector status, categorical dummy variables (sex, marital status, education and age) plus nine provincial dummies. The base group is non-unionized single male, with no post secondary education, age 30-39 working in British Columbia. Significance level: * p<0.10 ** p<0.05 *** p<0.01 in a two-tailed test.*

Table 4

Estimates of a regression model: logarithm of real hourly including time trend variable.

Explanatory Variables		
Union	0.131	(0.006)***
Time Trend	0.007	(0.003)***
Public	0.203	(0.003)***
Female	-0.226	(0.002)***
Married or Common Law	0.114	(0.002)***
Separated/Widowed/Divorced	0.071	(0.004)***
University Graduate	0.370	(0.003)***
College/some post secondary	0.161	(0.002)***
20 to 29 years	-0.199	(0.003)***
40 to 49 years	0.053	(0.003)***
50 to 64 years	0.031	(0.003)***
Newfoundland	-0.215	(0.004)***
Prince Edward Island	-0.254	(0.004)***
Nova Scotia	-0.184	(0.004)***
New Brunswick	-0.213	(0.003)***
Quebec	-0.086	(0.003)***
Ontario	-0.005	(0.003)***
Manitoba	-0.098	(0.003)***
Saskatchewan	-0.015	(0.003)***
Alberta	0.134	(0.003)***
Interactions of Union with:		
Time Trend	0.001	(0.005)***
Public	-0.021	(0.004)***
Female	0.062	(0.003)***
Married or Common Law	-0.041	(0.003)***
Separated/Widowed/Divorced	-0.018	(0.005)***
University Graduate	-0.067	(0.004)***
College/some post secondary	-0.027	(0.003)***
20 to 29 years	0.046	(0.004)***
40 to 49 years	-0.018	(0.004)***
50 to 64 years	0.008	(0.004)***
Newfoundland	0.093	(0.006)***
Prince Edward Island	0.095	(0.006)***
Nova Scotia	0.047	(0.005)***
New Brunswick	0.076	(0.005)***
Quebec	-0.003	(0.004)***
Ontario	0.014	(0.004)***
Manitoba	-0.023	(0.005)***
Saskatchewan	-0.029	(0.005)***
Alberta	-0.079	(0.005)***
Intercept	2.761	
R-squared	0.299	
Observations	689,129	

*Notes: Logarithm of Hourly Wage is the dependent variable. Survey weights are utilized. Standard Errors are presented in parenthesis. The explanatory variables are union status, time trend, public sector status, categorical dummy variables (sex, marital status, education and age) plus nine provincial dummies. The base group is non-unionized single male, with no post secondary education, age 30-39 working in British Columbia. Significance level: * p<0.10 ** p<0.05 *** p<0.01 in a two-tailed test.*

Table 5

Estimates of a regression model: logarithm of real hourly wage before and after the 2008 financial crisis.

Explanatory Variables	**Pre-2008**		**Post-2008**	
Union	0.127	(0.007)***	0.143	(0.008)***
Public	0.189	(0.004)***	0.215	(0.004)***
Female	-0.235	(0.002)***	-0.217	(0.003)***
Married or Common Law	0.113	(0.003)***	0.115	(0.003)***
Separated/Widowed/Divorced	0.069	(0.005)***	0.072	(0.005)***
University Graduate	0.387	(0.004)***	0.355	(0.004)***
College/some post secondary	0.166	(0.003)***	0.155	(0.003)***
20 to 29 years	-0.200	(0.004)***	-0.197	(0.004)***
40 to 49 years	0.051	(0.004)***	0.056	(0.004)***
50 to 64 years	0.021	(0.004)***	0.040	(0.004)***
Newfoundland	-0.275	(0.006)***	-0.156	(0.006)***
Prince Edward Island	-0.267	(0.006)***	-0.240	(0.006)***
Nova Scotia	-0.205	(0.005)***	-0.163	(0.005)***
New Brunswick	-0.232	(0.005)***	-0.195	(0.005)***
Quebec	-0.094	(0.004)***	-0.079	(0.004)***
Ontario	-0.003	(0.004)***	-0.009	(0.004)***
Manitoba	-0.108	(0.005)***	-0.088	(0.005)***
Saskatchewan	-0.052	(0.005)***	0.021	(0.005)***
Alberta	0.126	(0.004)***	0.141	(0.005)***
Interactions of Union with:				
Public	-0.015	(0.005)***	-0.027	(0.005)***
Female	0.068	(0.004)***	0.057	(0.004)***
Married or Common Law	-0.039	(0.005)***	-0.044	(0.005)***
Separated/Widowed/Divorced	-0.018	(0.008)***	-0.019	(0.008)***
University Graduate	-0.079	(0.005)***	-0.056	(0.005)***
College/some post secondary	-0.031	(0.004)***	-0.022	(0.004)***
20 to 29 years	0.050	(0.006)***	0.041	(0.006)***
40 to 49 years	-0.008	(0.005)***	-0.029	(0.005)***
50 to 64 years	0.032	(0.005)***	-0.015	(0.005)***
Newfoundland	0.111	(0.009)***	0.074	(0.009)***
Prince Edward Island	0.088	(0.009)***	0.101	(0.009)***
Nova Scotia	0.058	(0.008)***	0.108	(0.008)***
New Brunswick	0.084	(0.008)***	0.037	(0.008)***
Quebec	0.001	(0.006)***	0.069	(0.006)***
Ontario	0.006	(0.006)***	0.021	(0.006)***
Manitoba	-0.018	(0.007)***	-0.028	(0.007)***
Saskatchewan	-0.022	(0.007)***	-0.035	(0.007)***
Alberta	-0.098	(0.007)***	-0.061	(0.007)***
Intercept	2.785		2.797	
R-squared	0.303		0.295	
Observations	343,102		346,027	

*Notes: Logarithm of Hourly Wage is the dependent variable. Survey weights are utilized. Standard Errors are presented in parenthesis. The explanatory variables are union status, public sector status, categorical dummy variables (sex, marital status, education and age) plus nine provincial dummies. The base group for both regression is non-unionized single male, with no post secondary education, age 30-39 working in British Columbia. Significance level: * p<0.10 ** p<0.05 *** p<0.01 in a two-tailed test.*

Table 6

Regression results: estimates of the hourly real wage equations corrected for selectivity bias for pre and post credit crisis periods.

Explanatory Variables	**Pre-2008**		**Post-2008**	
Public	0.150	(0.002)***	0.164	(0.002)***
Female	-0.185	(0.005)***	-0.177	(0.007)***
Married or Common Law	0.019	(0.003)***	0.013	(0.003)***
University Graduate	0.346	(0.037)**	0.323	(0.026)**
College or some Post-Secondary	0.146	(0.028)**	0.134	(0.019)**
Age 20-29	-0.115	(0.017)**	-0.101	(0.011)**
Age 40-49	0.049	(0.017)**	0.025	(0.007)***
Age 50-65	0.058	(0.025)**	0.020	(0.013)**
Newfoundland	-0.162	(0.060)**	-0.084	(0.006)***
Prince Edward Island	-0.165	(0.007)***	-0.142	(0.007)***
Nova Scotia	-0.143	(0.006)***	-0.130	(0.006)***
New Brunswick	-0.142	(0.006)***	-0.128	(0.006)***
Quebec	-0.017	(0.004)***	-0.083	(0.004)***
Ontario	0.008	(0.004)***	0.019	(0.004)***
Manitoba	-0.119	(0.004)***	-0.121	(0.005)***
Saskatchewan	-0.070	(0.005)***	-0.019	(0.005)***
Alberta	0.022	(0.005)***	0.074	(0.005)***
Intercept	2.920	(0.193)	2.985	(0.133)
Number of observations	261,183		262,057	
Selection (Unionized)				
Female	0.041	(0.005)***	0.093	(0.005)***
Married or Common Law	0.001	(0.007)***	0.002	(0.007)***
University Graduate	0.365	(0.007)***	0.357	(0.007)***
College or some Post-Secondary	0.250	(0.006)***	0.258	(0.006)***
Age 20-29	-0.147	(0.009)***	-0.135	(0.009)***
Age 40-49	0.151	(0.007)***	0.091	(0.007)***
Age 50-65	0.231	(0.007)***	0.179	(0.007)***
Spouse-Unemployed	-0.051	(0.015)**	-0.083	(0.014)**
Intercept	-0.622	(0.010)**	-0.632	(0.010)**
Mills				
Lambda	0.041	(0.158)	-0.036	(0.100)
Rho	0.123		-0.105	
Sigma	0.339		0.344	

*Notes: Heckman two-step model. Selection procedure variables that affect the union selection plus an instrumental variable (spouse-unemployed) that is not included in wage equation. The first part of the table shows the results for the wage equations (excluding single participants) adjusted for selection bias. Standard errors are parentheses. Significance level: * p<0.10 ** p<0.05 *** p<0.01 in a two-tailed test.*