Luxembourg Income Study Working Paper Series

Working Paper No. 309

Unemployment Compensation Programs' Effect on the Employment of Young Men: A Cross-National Comparison of Canadian, British and American Unemployment Policies

Tess Heintze

July 2002

Luxembourg Income Study (LIS), asbl

Unemployment Compensation Programs' Effect on the Employment of Young Men: A Cross-National Comparison of Canadian, British and American Unemployment Policies

Tess Heintze
School of Social Work
Columbia University
th412@columbia.edu

July 2002

Unemployment Compensation Programs' Effect on the Employment of Young Men: A Cross-National Comparison of Canadian, British and American Unemployment Policies[1]

Abstract

This study compares the receipt rates of unemployment compensation of American young men (aged 18-25) with those from Canada and the United Kingdom. The results indicate that American young men are far less likely to receive compensation for their unemployment than are youth from the other two nations. Pooling the data from all three nations together we find that receipt of unemployment insurance in the prior year is complementary to current work participation whereas unemployment assistance benefits act as a work disincentive. This analysis indicates that the most effective policy alternatives are to increase the length of time that young men can accept insurance benefits and/or to raise the amount of income recipients can retain before the reduction of insurance benefits.

[1] The material from the 1995 U.K. Family Expenditure Survey is Crown Copyright; it has been made available by the Office for National Statistics through the ESRC Data Archive; and has been used by permission. Neither the Office for National Statistics nor the ESRC Data Archive bear any responsibility for the analysis or the interpretation of the data reported here.

I-I: Introduction

In the United States it has always been assumed that able-bodied men can and should support themselves and their households through employment. During the periods when men are not working, the American welfare state has been reluctant to offer any welfare benefit that may provide encouragement to be unemployed. To this end, means-tested transfer programs are not common in the U.S., particularly for males. Compensation for unemployment is generally provided through an insurance program only to those with a proven work history, those who are therefore most likely to return to workforce. It is not designed nor is it expected to compensate less experienced individuals for their unemployment. The only large-scale alternative program for men is the Food Stamps program.

This begs the question: how can men with little or no work history meet their basic needs legitimately if they face barriers or delays in gaining employment? This is particularly a problem for young men as they have only recently entered the workforce, or are entering it for the first time, and have had no opportunity to build up a lengthy record of employment. Their lack of work history is not an indication of their lack of interest in work, but rather of their youth. The job search process takes time and resources. By restricting access to these we may encourage youth to take the first job offered rather than the most suitable. Alternatively, we may discourage them from entering the labor force altogether, as they seek support through less legitimate or less acceptable means. By trying to ensure that welfare programs do not offer any work disincentives, and by not recognizing that gaining suitable employment may take time, we may have placed unreasonable constraints on the welfare of young American men. Little research has focused on the labor market response of young men to programmatic incentives, that is, the responses of those who are beginning their careers in the labor force. More work exists concerning teens and youth looking for casual employment and on older men. It is not clear that more liberal unemployment compensation programs would offer the same work disincentives to these young

men as they do to older men. Young men may require a more stable stream of income while unemployed to enable a more extensive job search. Older men, who in most instances have chosen their career paths and industries, may find work more easily and more quickly while unemployed. Older men may not require more liberal unemployment compensation. Thus the lack of availability of unemployment compensation for young men may not serve any positive function, but only serve to restrict their options.

How can we measure the disincentive effects of unemployment compensation on American young men when only a few have access to unemployment insurance and almost none to any unemployment cash assistance, or means-tested, program? One method is to pool together an American sample of young men with samples of young men who have better access to other insurance and assistance benefits. This will give some indication of the labor response of young men to unemployment compensation programs, even if it does not isolate the response of American young men. To this end, this study pools together samples of young men aged 18 to 25 from Canada, the United Kingdom and the United States. It assesses their labor responses to receipt of unemployment insurance and assistance. Section I-II offers an overview of current research into the disincentive effects of unemployment assistance and unemployment insurance, an exploration of the selection of these three nations in particular and a brief review of the regulations of their unemployment compensation programs. Section I-III contains a more formal presentation of the research questions. The guiding theory is presented in Section I-IV, with the econometric specification in Section I-V. Results, the discussion and conclusion follow in Sections I-VI and I-VII. Throughout this analysis 'unemployment compensation' is used as a general term to cover both unemployment insurance schemes (or contributory programs) and unemployment assistance (means-tested programs).

I-II: Background

(a) Unemployment Compensation Programs and their Effects on Work Behavior

In western democracies it is common for unemployment to be compensated, but different nations choose different avenues to achieve this goal. Some offer only unemployment insurance, others only unemployment assistance and others both programs. The two types of program are however very different, and to confuse them is to lose some of the detail of the effects of unemployment compensation programs on work behavior (Atkinson & Micklewright, 1991).

Unemployment insurance programs are more familiar to most Americans than unemployment assistance. Insurance programs require contributions to be paid into a fund and when the insured individual experiences unemployment a time-limited benefit is paid to him. The employer, the employee or both may pay these contributions. Unemployment insurance benefits are usually only available to those who have worked for the period of time required to vest themselves in the scheme. It is generally not available to those who are new entrants to the labor force, have only held a job for a limited length of time, those who leave work voluntarily without good cause or those discharged for misconduct related with the work. Its aim is to provide temporary income support to experienced workers who are unemployed through no fault of their own (Wandner & Stengle, 1997). It is not supposed to compensate all workers for their unemployment. In the U.K. and the U.S. between 30 and 40 percent of the unemployed receive insurance benefits (Atkinson & Micklewright, 1991; Wandner & Stengle, 1997). The young are the least likely to receive benefits (Atkinson & Micklewright, 1991).

Unemployment assistance programs, on the other hand, offer benefits to all unemployed whose income falls below a prescribed level, that is, they are means tested. The means test is usually applied to the income of the entire household in which the unemployed individual resides. Often separate benefits levels and means tests are required for youth, particularly those under the age of

majority. Some countries do not offer an assistance program that is specifically for the unemployed, but potentially employable non-working individuals can apply for benefits under a general, broader social assistance scheme (Atkinson & Micklewright, 1991). Benefits tend to be lower in an assistance scheme than in an insurance program, but as long as the household income is low enough, they are generally not time-limited[2].

Overall, insurance benefits are held to create a disincentive to work, with higher benefits creating longer unemployment spells (Lancaster & Nickell, 1980; Danziger, Haveman & Plotnick, 1981; Moffitt & Nicholson, 1982; Meyer, 1990). This is found to be the result both of more generous benefits and of the higher disposable income resulting from more favorable tax treatment of benefits (Anderson & Meyer, 1997). Research on the aggregate level, rather than the individual level, indicates that higher benefits push up real wages and thus increases unemployment (Pissarides, 1991). However, the effect of benefits on unemployment is seen to vary with assumptions about the benefit structure, and with the age of the recipients (Atkinson & Micklewright, 1991).

Benefit duration also appears to have an effect on labor market behavior. Many researchers have found that insurance recipients' return to work is highly associated with the length of time that the individual expects to retain eligibility for benefits, creating a 'spike' in the hazard of returning to work that appears just before benefit exhaustion (Mortensen, 1977; Moffitt & Nicholson, 1982; Milbourne, Purvis & Scoones, 1991). Meyer (1990), however, indicates that the individual cannot always predict benefit duration accurately and that few spells last long enough to be effected by the spike. He thus indicates that this return to work (just before the cessation of payments) may not be solely the effect of approaching benefit exhaustion. Meyer maintains that

[2] Time limits on collection of means-tested benefits were introduced in the United States with the implementation of the 1996 Personal Responsibility Work Opportunity and Reconciliation Act. The data in

benefit level is much more influential than benefit duration on the length of unemployment (Meyer, 1990).

Unemployment insurance can have some unintended effects other than increasing the length of unemployment. It may be used to fund temporary layoffs, particularly in the United States where temporary layoff has been seen as the reason for half of all unemployment (Atkinson & Micklewright, 1991). However, few individuals actually return to the same employer. Green and Sargent (1998) see the tailoring of job duration (as opposed to unemployment duration) to unemployment insurance vesting periods as much more a feature of seasonal employment than regular employment in Canada. They find, however, that this 'gaming' effect is too small to justify much concern about altering the length of the benefit qualification period.

The work disincentive effects of unemployment assistance are, in theory, more clear-cut. The static labor model implies that assistance programs reduce work effort. An assistance benefit is reduced as the individual works and indeed is often reduced dollar for dollar after a small disregard in income. This implies that the individual may face an implicit tax rate on benefits of more than 100 percent, as he may pay income tax on the earnings as well as lose a benefit dollar for every gross earned dollar[3]. Therefore the income and the substitution effects both act in the same direction, reducing work effort (Moffitt, 1992).

As cash unemployment assistance is not available to all men in the United States, its effects on American men have not been tested. A study done in Michigan after the General Assistance program was dismantled indicate that this program was used by some men as a form of unemployment assistance (Kost, 1996). However, the men in the sample generally had very low

---

this analysis predates this legislation.

[3] Note that the Earned Income Tax Credit offsets some of this disincentive for Americans.

levels of human capital and were thus highly likely to be unemployed with or without General Assistance. To assess the impact that the program alone had on their labor behavior was difficult to calculate. Furthermore, the sample was small with the study taking the form of qualitative interviews rather than quantitative analysis.

One national means-tested welfare program that is available to men is the food stamp program. Although the program does not provide cash benefits it is often considered to be a near-cash alternative. The majority of families spend more on food than is covered by the value of the food stamps. On average, an eligible household receives $525 per quarter in food stamps and spends $631 per quarter on food (U.S. House of Representatives, 1996; Cage, 1994). This indicates that food stamps act as a cash supplement to the household as, if cash were to be provided and the buying habits of the household do not change, the provision of a cash benefit in the same amount would merely supplement the purchase of food[4]. Studies indicate that, except in the poorest of households, the cash equivalent of food stamps is more than 90 percent of their value (Fraker & Moffitt, 1988). Despite consideration of food stamps as a form of near-cash social assistance, few studies examine the effects of food stamp receipt on labor behavior. Fraker and Moffitt (1988) examine the joint impact of Aid to Families with Dependent Children and Food Stamps on the labor supply of women. Although a different population of interest from that considered here, food stamps have an effect similar to that of other assistance programs in that it reduces the labor supply of participants by about nine percent (Fraker & Moffitt, 1988).

Concern about the unintended effects of unemployment compensation should not blind us to the positive benefits offered by the programs. When workers are unemployed, unemployment compensation continues a flow of income into the household allowing is members to continue to

[4] This is, of course, a simplification as the provision of a cash benefit as opposed to food vouchers may well reduce the amount of food bought.

meet basic needs, particularly in the case where the unemployed individual was the only working member of the household. Where the benefit levels are 'optimally' set, research indicates that unemployment compensation, particularly unemployment insurance, can have a positive effect on welfare (Hammermesh, 1982; Hansen & Imrohoroglu, 1992). Additionally, unemployment compensation does fund the job search process; individuals who receive benefits undertake a more thorough search for employment while out of work, increasing their chances of employment (Tannery, 1983; Ben-Harim & Suckerman, 1987; Wadsworth, 1991). Although this search effect may be outweighed by an increase in the reservation wage that may arise from the presence of that unemployment benefits, it offers an employment incentive rather than disincentive. Research indicates that the effect that unemployment compensation may have on the job search process is larger where there are liquidity constraints among the unemployed (Wadsworth, 1991) as is more likely to be true for young men.

Overall, prior research indicates that unemployment benefits will have a negative effect on work behavior. However, little if any of this body of research has been done on young men. Young men do face a unique situation in the job market. Young men are just beginning their working lives. While they are most likely to be unemployed, they are least likely to receive benefits, particularly in the United States (Atkinson & Micklewright, 1991). These situations are different from those of older men and imply that unemployment compensation may have different effects on their work behavior than on the behavior of older men or women.

(b) Nations Chosen for Comparison

To analyze the effect of different unemployment compensation programs on work effort, we need to compare the reactions to unemployment compensation across different nations. We cannot compare the effects of cash unemployment assistance, for example, using just American data because the program does not exist in the U.S.. Although unemployment insurance programs do

vary across the country, variation is not as strongly marked as it is across nations. For example, duration of unemployment benefits does not vary greatly across the American states. Furthermore, few young men in America receive unemployment insurance making it difficult to assess the effect insurance has on the labor behavior of young men. In undertaking cross-national comparison however, we need to ensure that we compare across countries that are as similar as possible. This provides some control for different national values, approaches to welfare in general and economic factors. In other words, as much as is practicable, the analyzed variation in the work effort of young men should be the result of particular unemployment compensation policies rather than other national factors.

The three countries studied here, Canada, the U.K. and the U.S., are all 'Anglo' countries. They have a similarity in culture, values, religion and language that influences their economic and political choices (Lipset, 1989; Gornick & Jacobs, 1994). Their histories are intertwined in many aspects as both people and ideas flow across the Atlantic. Changes in public policy, in particular welfare policy, and some alterations in ideology, that occur in one country often reappear in the others, albeit in a somewhat altered aspect (Trattner, 1994; Savoie, 1994). The differences between Anglo countries are smaller than those between the English-speaking countries and other nations (Lipset, 1989).

The three Anglo countries included in this analysis are also associated through the characteristics of their welfare states. The principle typology of welfare states is that formulated by Esping-Anderson (1990). He classifies three forms of welfare state: the liberal, social democratic and corporatist regimes. Esping-Anderson identifies one factor upon which to distinguish between the different welfare states as the method the nation uses to overcome Baumol's cost disease (Baumol, 1967) as well as the extent to which the country manages to do so (Kloosterman, 1994). Baumol's cost disease assumes that productivity in the services sector will lag behind that of

manufacturing. If wages rise and are linked between sectors, the increased cost of labor in the manufacturing sector can be paid out of increased productivity, but this is not the case in the services sector. Here prices may rise to cover labor costs and these products can thus price themselves out of the market as household chose to produce more services for themselves. In such cases, the services sector will stagnate. Esping-Anderson also considers the decommodification of labor in his analysis. This is the extent to which individuals must sell their labor as a commodity on the market in order to survive, as opposed to having a right to welfare support. Nations with more generous welfare support offer an economic environment where it is less of a necessity for individuals to sell their labor, that is, labor is less of a commodity (Esping-Anderson, 1990).

In Esping-Anderson's liberal regime, Baumol's cost disease is contained by low wages in the services sector, which tends to be unprotected by unions. Wages are not linked between the services and manufacturing sectors. Services such as child-care and house cleaning can thus be bought on the market at low cost, freeing women to enter the labor market. Decommodification is low, that is there are few rights to social welfare. Esping-Anderson (1990) identified the U.S. and Canada as clearly belonging to the liberal regime. The U.K. has somewhat mixed programs, but is also identified with the liberal welfare state. In the social democratic regime, Baumol's cost disease is overcome by services such as childcare being provided by the government. Decommodification and welfare rights are high. The welfare state in these nations expands with much of the resulting growth in the public sector workforce being among females. Sweden is seen as the archetypal country of this approach to welfare. In the final type of welfare state, the corporatist regime, neither approach is taken and Baumol's cost disease is unassauged. Germany is offered as the prime example. Here welfare rights are attached to class and status. According to this theory then, one method of typing welfare states is by the size of the services sector, the

participation of women in the labor force and whether child-orientated services are offered by government (Kloosterman, 1994).

Table I-1 offers these statistics for the nations of the study and for the other two regimes archetypes: Sweden and Germany. The percentage of the workforce in the services sector is lowest in Germany, with strong similarities between Canada, the U.K. and the U.S.. The percentage of the labor force that is female is also lowest in Germany. Government child policies are assessed using the work of Gornick, Meyers and Ross (1997) who created an index to measure generosity of government child-related policies and applied it to 14 western democracies. Both index values and the ranking (in parenthesis) are provided in table I-1. As expected by the Esping-Anderson typologies, the Swedish welfare system is markedly more generous than any of the other nations shown. The U.K. and the U.S. are clearly the least generous. According to Esping-Anderson such services are provided in the private sector in these nations. Overall there is a similarity between Canada, the U.K. and the U.S. in the service sector, female labor force participation and public child policies indicating some support for placing them in the same Esping-Anderson welfare state typology. In turn this implies the welfare systems of the three countries display strong similarities, especially when compared to nations that fall into the other Esping-Anderson categories.

< Table I-1 about here >

With specific regard to the unemployment compensation systems of the three countries, they are similar in that, compared to other western nations, their programs are among the least generous. In a 1991 OECD ranking of unemployment compensation benefit generosity in nineteen countries (all western economies with the exception of Japan) Canada, the U.K. and the U.S. all rank in the lower half of the distribution. Canada ties for eleventh position with Germany, the U.K. is

sixteenth and the U.S. seventeenth. Of the four countries that rank between Canada and the U.S., two are the other Anglo countries included – Australia and New Zealand – confirming the similarity between English-speaking nations as compared to other nations (OECD, 1994; Gornick, 1998).

Turning from similarities in politics and public policy to the economy, indicators of the similarity of the three nations' economies are provided in table I-2. At the time of analysis (1995 for the U.K. and 1994 for the U.S. and Canada), all three economies were in expansion (OECD, 1998). In all three countries approximately 85 percent of men were in the labor force. Unemployment was comparable in Canada and the U.K. but much lower in the U.S. Unemployment is measured using the 1982 International Labor Organization (ILO) Guidelines. This implies that the unemployed are those who are without work who are actively seeking a job.

< Table I-2 about here>

Table I-2 also contains poverty and inequality measures. Poverty is much more prevalent in the U.S. than in Canada or the U.K. The measure used here indicates the number of households who have equivalent income under 50 percent of the median national equivalent household income. Although some argue for absolute scales, having an income below fifty percent of the median income is a widely accepted measure of poverty (Pressman, 1995; Smeeding, 1995). The equivalence scale used to control for household size is that proposed by the OECD (1982)[5]. The income used is the disposable household income, that is, after tax and transfer income. All monetary amounts used have been converted to the equivalent of 1995 American dollars using purchasing power parities. Purchasing power parities compare the cost of basic foods across

nations. Although these have been criticized because not all cultures eat the same foods (Pressman, 1995), the nations being compared here are similar enough in culture and food tastes for this to be less of an issue.

The last measure included on table I-2 is a measure of inequality. It is the ratio of the equivalent disposable household income at the 90th percentile to that at the 10th. It shows that, for example, wealthy Canadian households enjoy nearly four times as much disposable income as their poor compatriots. Clearly America displays far greater inequality than does either the U.K. or Canada.

Given these measures of poverty and inequality, it may appear that the U.S. is not particularly comparable to Britain or Canada in this regard. While this cannot be disputed, other nations have far lower poverty rates and inequality measures (Sullivan & Smeeding, 1997; Korpi & Palme, 1998). When comparing OECD countries, these three nations tend to cluster (together with the other Anglo nations) at the high end of the distribution of measures of poverty and inequality. Furthermore, while the U.S. suffered a large rise in inequality during the 1980s, that experienced by the U.K. was even greater. Canada, on the other hand, had little or no change in inequality during the 1980s and early 1990s (Smeeding, 1995; Smeeding & Gottschalk, 1996).

All in all, three countries chosen for analysis are remarkably comparable in welfare state typology, male employment and poverty. One cannot expect different countries to be identical (which these certainly are not) but the three display more similarities than differences, especially when compared to other nations.

[5] The divisor is calculated as follows: {1 + (0.5*Number of Children) + (0.7* Number of Adults)}/2.2 This implies that it is 1 for a household containing one child and one adult, and each additional child adds 0.228 to the scale and each adult, 0.318.

(c) Current Unemployment Compensation Policies:

As these countries are not identical, we need to assess specifically how their unemployment compensation programs vary. I offer a brief overview of unemployment insurance in the three nations, followed by a short examination of the unemployment assistance programs.

(i) Unemployment Insurance

Unemployment insurance is an existing program in Canada, the U.K. and the U.S.. In all three of these countries unemployment benefits are predicated on contributions to the social security system made by both employers and employees and thus receipt depends on previous employment. In Canada and the U.S. benefits are a percentage of earnings up to a prescribed maximum; in the U.K. benefits are set at a fixed amount. The U.S. system varies by state, whereas in Canada and the U.K. the system is centrally administered and is uniform across the nation.

In order to qualify for benefits in Canada a full-time employee must have worked for between 12 and 20 weeks during the previous year. The actual number of weeks depends on the local unemployment rate: if this is 6 percent or less, 20 weeks of work are needed for qualification; if 13 percent or more, 12 weeks are needed. With each increase of 1 percent in local unemployment, one week less work is required for benefit eligibility. New entrants to the system are required to work for 26 weeks before receiving benefits. Individuals who earn less than 20 percent of the maximum insurable earnings do not pay unemployment contributions and are thus not eligible for benefits. Benefit recipients are paid 55 percent of their average gross earnings up to the maximum insurable earnings level. In 1994 (the year of analysis) the maximum insurable earnings were C$815, and therefore the maximum benefit was C$448. Benefits will be paid for between 14 and 45 weeks, depending both on the individual's weeks of insurable employment, and the local unemployment rate. An individual can earn up to 25 percent of his weekly benefit

without loss but earnings above this are deducted from the benefit on a dollar for dollar basis (Corak & Jones, 1995; OCED, 1996; HRDC, 1996).

In the U.K. benefits are predicated not on time in the workforce per se, but rather the amount of contributions made to the system personally or on their behalf; contributions can also be credited to the individual during periods of training or unemployment. Individuals are eligible for benefits if they have paid at least 25 times the weekly lower limit earnings in one of the past two tax years and have also paid or been credited with at least 50 times this limit in each of the past two years. In 1995 the weekly lower limit was £58. At that time contributions were paid by, and on behalf of, employees who earned less than this weekly lower limit; however as of 1999 such individuals and their employers are exempt from payment of contributions. Benefits in 1995 were £46.45 per week for one person, with an additional £28.65 for a partner earning less than this additional allowance. In 1999, benefits increased to £80.65 for a couple, and £51.40 for a single individual over the age of 25. The unemployment benefit is payable for 6 months and is available not only to the insured unemployed but also those working 16 or fewer hours per week. If the individual work and collects benefits, the first £5 of earnings per week are disregarded (OECD, 1996; Inland Revenue, 1999).

The unemployment insurance system is less easy to describe in the U.S. as it is administered at the state level and thus varies across the country. In general, an individual must have earned a particular amount during a base period prior to unemployment. This base period is usually the first 4 of the last 5 calendar quarters worked. Monetary eligibility requires earnings at a particular level during the quarter in the base period with the highest earnings, with total earnings during the base period not falling below a multiple (usually between 1.25 and 1.5) of this highest earnings quarter. Generally, the individual must earn between US$600 to US$2,500 in this highest earnings quarter. Weekly benefits in the majority of states are between 1/20 and 1/26 of

the earnings in the highest quarter. Other states[6] compute percentages of the weekly wage (the percentage being between 50 and 67), and the remainder of the states calculate a multiple of the annual wage[7]. All states issue a maximum and minimum benefit level with 15 states offering a dependent's allowance. Maximum weekly benefits vary from US$190 to US$400. In general, benefits may be received for up to 26 weeks[8], 30 in Massachusetts and Washington State. All states allow combinations of work and benefit receipt. Most have a small disregard of earnings with a further loss of benefit on a dollar for dollar basis. Five states reduce earnings on a less than dollar for dollar basis: Montana, Nevada, Washington, California and Alaska (Anderson & Meyer, 1997; OECD, 1996; U.S. Department of Labor, 1999).

(ii) Unemployment Assistance

Assistance programs are those where eligibility is determined by a means test rather than by prior contributions to an insurance fund. Only the U.K. has a true unemployment assistance program, a program targeted specifically on the unemployed, but Canada has social cash assistance programs aimed at the low income for which low income unemployed are also eligible.

Canada's social assistance program is administered at the provincial level. Although varying widely from province to province, each of the twelve programs shares certain aspects. Each requires a needs test (where budgetary needs of the household are compared with household assets) and generally a very small level of liquid assets are exempt from consideration. Each province and territory uses a different method of calculating the basic social needs that the benefit is designed to cover. Benefits for single employable adults vary across the provinces from C$48 per week in Newfoundland to C$116.27 in British Columbia. The territories have higher rates.

[6] Minnesota, New Jersey, New York and Ohio
[7] Alaska, Kentucky, Montana, New Hampshire, Oregon, and West Virginia

Again these payments are made as long as income and assets are low enough for eligibility. Work and welfare may be combined, with most provinces allowing a loss of less than a dollar of benefits for each dollar earned (OECD, 1996; National Council of Welfare, 1998).

Like the Canadian provinces, the U.K. offers an assistance program that pays the difference between individuals' means and a set amount to cover basic needs. The applicable amount for basic needs is similar to that offered under the unemployment insurance program, with a slightly lower maximum benefit for those aged between 18 and 24. The disregard for the assistance program is also £5, £10 for couples (OECD, 1996).

The U.S. does not have a major social cash assistance program for which unemployed men are eligible. Aid to Families with Dependent Children (AFDC), the welfare program in place at the time of data collection, included a program for unemployed parents in each state, but only 6% of AFDC recipient families in 1994 received aid under this program (approximately 2/5 of 1 percent of the general population) (U.S. House of Representatives, 1996). It is thus not widely available. Some localities offer general assistance to men but duration and amount of benefits are so restricted, and the number of programs in existence around the country are so few, that the effects of this program are here disregarded. The largest national assistance program available to men is the Food Stamps program. While this program does not provide cash benefits, it does provide vouchers for food. As discussed above, most eligible families buy at least as much food as the vouchers cover, the program acts as a near cash alternative to a true social assistance program.

The food stamps program is funded federally, but day-to day administration is carried out by the states. Benefit calculation is complex, but the benefit size is dependent on the household's size,

[8] Benefits in Alaska, California, New Jersey, Connecticut, Oregon and the District of Columbia are extended when state unemployment reaches specified levels. Weeks of coverage are also extended in

net monthly income and maximum monthly benefit levels. Once the household's net income is established (and thus the maximum benefit), the actual benefit is attained by subtracting an amount that the household is expected to contribute toward food purchases (generally 30 percent of its income). In 1995, monthly benefits average $71 per person and $175 per household. Those who are on strike, illegal and temporarily resident aliens[9] and most post-secondary students are also not eligible for benefits (U.S. House of Representatives, 1996).

I-III: Research Questions

This analysis is concerned with the labor response to the receipt of unemployment compensation, both assistance and insurance benefits, by young men. As the United States has no cash assistance program for men, the best alternative, the Food Stamp program, is used to measure the effects on assistance receipt on the labor response of young American men. The young men in this sample are aged between 18 and 25 and are in the labor force. They reside in Canada, the United Kingdom and the United States. Those who are still pursuing education are not included unless they have a job or declare themselves to be actively seeking work. However, those who are in school and who either have employment or who are actively seeking it are included in the analysis.

I have in the above discussion assumed that young Americans are less likely than Canadian or British men to receive unemployment compensation. Additionally I have asserted that American young men have a smaller chance of gaining unemployment compensation than their older counterparts. This will be explored during this analysis. I therefore ask: are young American

---

Hawaii when a man-made or natural disaster occurs.

[9] This category for non-eligibility was broadened as part of the 1996 Personal Responsibility Work Opportunity and Reconciliation Act but, as mentioned earlier, the data in this analysis predates this legislation.

men less likely to receive compensation while unemployed than their British and Canadian counterparts? Are American young men less likely to receive benefits while unemployed than their more senior compatriots?

I am also interested in the labor response of young men to unemployment compensation and the effects of unemployment compensation program design on employment decisions. This response will be assessed as a joint decision between work and receipt of unemployment compensation, both unemployment insurance and unemployment assistance. Due to data restrictions, I examine the effect of unemployment compensation receipt in the past year on current work participation. I ask: do the maximum benefits available last year affect a young man's decision to work? How does the structure of the programs affect the employment of young men?

I-VI: Theory

Previous research indicates that the receipt of welfare benefits (from both assistance and insurance programs) does affect work behavior. Past program participation is strongly related to current program behavior. If we wish to examine the labor market behavior of young men, we must therefore also consider their participation in welfare programs. Joint estimation of participation in work, assistance and insurance programs will account for the bivariate selection issues that would arise if we analyze work behavior by itself rather than in conjunction with both unemployment compensation programs.

However, for simplification purposes, let us begin by assuming we are analyzing only the work behavior of young men. In general, labor supply can be modeled as follows:

$$H_j = \beta_H X_j + \varepsilon_{Hj} \qquad \text{if } \beta_H X_j + \varepsilon_{Hj} > 0 \Leftrightarrow Y_{Hj} = 1$$

$$H_j = 0 \qquad \text{if } \beta_H X_j + \varepsilon_{Hj} = 0 \iff Y_{Hj} = 0 \qquad \text{(I-1)}$$

where H_j is the hours worked by individual j, X_j is a vector of exogenous variables that affect a individual's decision to work, β_H is a vector of unknown parameters and Y_{Hj} is a dummy variable indicating whether the individual works. The disturbance ε_{Hj} is a random variable that indicates the "taste" for time in market work versus time in other, nonmarket activities. It is assumed to be independently and normally distributed with mean zero. This formulation of the labor supply model allows for young men who do not enter market work ($Y_{Hj} = 0$), and who thus have no hours of work ($H_j = 0$), that is, young men whose taste for work is comparatively low ($\varepsilon_{Hj} < -\beta_H X_j$). As we wish to examine young men's decisions to enter employment, rather than their decisions concerning the number of hours to work, we need to explore the probability that their taste for work is comparatively high, that is the probability that $\varepsilon_{Hj} > -\beta_H X_j$. When this occurs, $H_j > 0$ and the young man works ($Y_{Hj} = 1$).

When we expand the model to include young men's participation in both welfare programs we obtain three equations:

$$\begin{aligned}
H_j &= \beta_H X_j + \varepsilon_{Hj}, & Y_{Hj} &= 1 \text{ if } \varepsilon_{Hj} > -\beta_H X_j, & Y_{Hj} &= 0 \text{ otherwise} \\
P^*_{Ai} &= \beta_A X_j + \varepsilon_{Aj}, & P_{Aj} &= 1 \text{ if } \varepsilon_{Aj} > -\beta_A X_j, & P_{Aj} &= 0 \text{ otherwise} \\
P^*_{Ij} &= \beta_I X_j + \varepsilon_{Ij}, & P_{Ij} &= 1 \text{ if } \varepsilon_{Ij} > -\beta_I X_j, & P_{Ij} &= 0 \text{ otherwise}
\end{aligned} \qquad \text{(I-2)}$$

Here the H subscript refers to work, the A to unemployment assistance programs and the I to unemployment insurance programs. P^* is a latent variable that measures the probability that individual j will participate in an unemployment compensation program, P is a dummy variable indicating actual participation in either program, X is a vector of exogenous variables that affect

the individual's decision to work, or to participate in either unemployment compensation program and β_H, β_A and β_I are vectors of unknown parameters. The errors ε_{Hj}, ε_{Aj}, ε_{Ij} are trivariate normally (TVN) distributed. All have means of zero and are correlated ρ_{HA}, $\rho_{HI,}$ and ρ_{AI}.

Equations (I-2) assume that all individuals j are eligible for both forms of unemployment compensation. In all three countries (Canada, the U.K. and the U.S.) young men can make themselves eligible for assistance payments (note that we are concerned only with those over the age of 18) either by claiming their portion of the household's assistance payment as one of the adults in the household or, if their family or household has too much non-labor income to receive assistance, by forming into their own household. It is possible that young men will hold assets in their own right that make them ineligible for assistance (that is, their means are too high to be eligible), but in this given age group (18-25) it is unlikely (Hatcher, 1999). We will thus assume that young men can chose to adjust their labor response and living arrangements to make themselves eligible for assistance payments if they wish to do so.

Eligibility for unemployment insurance, however is dependent on past labor experience and involuntary job loss. As indicated above, young men are less likely than older men to be eligible for unemployment insurance payments. Thus we adapt our model (I-2) to reflect this:

If $E_I = 1$:

$$H_j = \beta_H X_j + \varepsilon_{Hj}, \qquad Y_{Hj} = 1 \text{ if } \varepsilon_{Hj} > -\beta_H X_j, \qquad Y_{Hj} = 0 \text{ otherwise}$$

$$P^*_{Ai} = \beta_A X_j + \varepsilon_{Aj}, \qquad P_{Aj} = 1 \text{ if } \varepsilon_{Aj} > -\beta_A X_j, \qquad P_{Aj} = 0 \text{ otherwise}$$

$$P^*_{Ij} = \beta_I X_j + \varepsilon_{Ij}, \qquad P_{Ij} = 1 \text{ if } \varepsilon_{Ij} > -\beta_I X_j, \qquad P_{Ij} = 0 \text{ otherwise}$$

$$(\varepsilon_{Hj}, \varepsilon_{Aj}, \varepsilon_{Ij}) \sim \text{TVN}(0, 0, 0, 1, 1, 1, \rho_{HA}, \rho_{HI,} \rho_{AI})$$

If $E_I = 0$:

$$H_j = \beta_H X_j + \varepsilon_{Hj}, \qquad Y_{Hj} = 1 \text{ if } \varepsilon_{Hj} > -\beta_H X_j, \; Y_{Hj} = 0 \text{ otherwise}$$

$$P^*_{Ai} = \beta_A X_j + \varepsilon_{Aj}, \qquad P_{Aj} = 1 \text{ if } \varepsilon_{Aj} > -\beta_A X_j, \quad P_{Aj} = 0 \text{ otherwise}$$

$$(\varepsilon_{Hj}, \varepsilon_{Aj}, \varepsilon_{Ij}) \sim \text{BVN}(0, 0, 1, 1, \rho_{HA}) \qquad \text{(I-3)}$$

where E_I is a dummy indicating eligibility for unemployment insurance, TVN for trivariate normal and BVN, bivariate normal.

I-V: Data and Econometric Specification

(a) Data

The data used in this analysis comes from Wave IV of the Luxembourg Income Study (LIS). LIS offers a database containing income survey data from multiple nations. LIS incorporates microdata that are currently in existence and manipulates the data such that the variables are comparable cross-nationally. Note that although the resulting variables are similar between datasets, they cannot be identical. Total comparability across different political systems, value systems and cultures is only theoretically possible. However, we can assume, for example, that a variable measuring total household income captures the same sources of income across nations. In other words, as Smeeding (1995) indicates, we can assume that comparability is of an acceptably high standard.

The three national datasets used here contain data from the 1994 Canadian Survey of Consumer Finances, the 1995 U.K. Family Expenditure Survey and the 1994 U.S. March Current Population Survey. As the study examines the behavior of young men, the subset of all men between the ages of 18 and 25 inclusive who are in the labor force (that is, those employed or actively seeking work) is retained for analysis. The final sample includes 3,626 Canadians, 536 Britons and 5,793

Americans (unweighted N = 9,955). Weights are included to make individual data representative of the national population of young men in the labor force. Males between the ages of 18 and 25 comprise 6 percent of the Canadian population, 5 percent of the British population and 6 percent of the American population.

Means (and standard deviations where applicable) of the variables included in this analysis are displayed in table I-3. The demographic variables included are for the most part self-explanatory. The only variables that require description are the education variables. The three national datasets included in the analysis measure educational attainment in very different ways. The British dataset measures only age at which formal education ceased, while the Canadian and American education data is offered in a format with 7 to 8 mutually exclusive categories. The four levels of education included here approximate less than a high school education, high school graduation, some college or an associate degree and bachelor's degree and above[10]. Within this understanding of the variable, the educational attainment for young men differs across the three countries. British young men are less likely to have finished high school, educational level 1, but the education system in Britain begins a year earlier than that of the U.S. and also has an extra year of high school (the equivalent of two 'senior' years). Nationally recognized academic attainment examinations are completed after the fifth year of high school (similar to American junior year) in Britain and thus not completing the two-year sixth level of high school may not have similar consequences to not attaining a high school diploma in the U.S.. In other words, those in education level 1 in Britain may be more 'educated' and better able to participate in the workforce than their educational counterparts in Canada and the U.S. Thus the high percentage of British young men in education level 1 (nearly fifty percent) as compared to Canada and the U.S. (twenty-five and twenty-two percent respectively) may be of less concern than first appears.

Similarly, the higher percentage of American young men with a tertiary degree, as compared to Britons and Canadians, may not reflect that Americans are more likely to be better educated than Canadians or the British. To summarize, these measures are not exact and it is best to conceive of educational attainment as very low, low, medium and high rather than as specific levels of attainment. The education variables more precisely measure young men's educational attainment as compared with their compatriots rather than with their cross-national counterparts.

<Table I-3 about here>

Nearly all of the other demographic variables are similar cross-nationally, one exception being the lower marriage rate of Canadians. They are also less likely to form their own households. The two phenomena are, most likely, associated. Unfortunately, due to the formulation of the dataset, we cannot assess whether children in the household are offspring of the young men who are the target of this study or whether the young men live with the children of other individuals. We only have a count of children in the household. Thus all analyses include a variable that measures whether the young man resides with his family of origin (that is, with any relative excepting his wife) or in his own household. This latter category includes young men who live alone, with friends or with their spouse. The other demographic variable that is dissimilar cross nationally is housing tenure. Britons have, on average, a far longer housing tenure.

The income variables include a measure of other equivalent income. This measure assesses that portion of the household gross income that is not contributed by the young men being studied. This 'other income' is then divided by an equivalence scale to control for the size and make-up of the household. As discussed above, the scale used here is that proposed by the OECD (1982).

[10] In the U.K. bachelor's degrees are three year degrees rather than four year degrees as in Canada and the U.S. Additionally in the U.K. there is an extra year of education at the end of high school, the equivalent of

All monetary amounts used have been converted to the equivalent of 1995 American dollars using purchasing power parities discussed above. The maximum unemployment insurance and assistance benefit variables reflect the previous discussion about the various unemployment compensation programs. A variable is included to control for the maximum length of time unemployment insurance benefits can be claimed. No equivalent variable is included for unemployment assistance as the program was not time limited during 1994-95 in any of the included countries. Note that unemployment assistance payments are generally less generous than those paid by unemployment insurance programs except in Britain where they are the same. The final programmatic variable included in the analysis measures the amount of unemployment compensation that is disregarded when the benefit is reduced due to labor income. This variable, disregard, is included for both the unemployment assistance and the unemployment insurance programs.

The final set of variables describes the residence and regional labor market of individuals. The population density variable indicates that the sample member does not live in a major urban or suburban area but rather in a rural area, a small city or town. The unemployment measure included is that for each state in the U.S., each province in Canada and each Government Office Region in the U.K. The unemployment rates are all calculated according to 1982 International Labor Organization (ILO) Guidelines. This implies that the unemployed are those who are without work who are actively seeking a job. Discouraged workers are not included nor those not in the labor force. Note that this definition differs from that popular in the U.K.; British unemployment measures often reflect the percentage of the population claiming benefits rather than that conforming to ILO guidelines.

---

a lower and upper senior year in additional to the freshman, sophomore and junior years.

(b) Econometric Specification

We will begin by comparing the labor participation, wages and unemployment compensation of men cross-nationally to assess the national characteristics that underlie the heart of our analysis. Following this, two labor market models will be estimated using the system of equations outlined above (equations I-3). The first model examines the effects of the program variables maximum weekly assistance benefit and maximum weekly insurance benefit, together with the maximum duration of benefits. The second model adds the effects of the assistance and insurance benefit disregard of labor income. All demographic, income and regional variables are included in both models. Due to the structure of the LIS dataset, both models are examined using the young man's work status at the current time, together with dummies indicating whether or not he participated in unemployment assistance and insurance programs in the past year.

However in order to implement these model three econometric issues require to be solved. The first is the imputation of missing, but necessary, data; the second is implementing the trivariate probit (that is, the calculation of trivariate normals); and third, we need to consider how to calculate a predicted wage for all young men in the data set and to perform the necessary recalculations of the standard errors.

(i) Imputation of Missing Data

The model is dependent on the insurance eligibility of the young man. Unfortunately the LIS dataset does not contain past labor information that would enable us to assess whether the young man is eligible for insurance payments or not. Equally, we cannot know whether the individual is eligible because he left work involuntarily but not through misconduct. However, for some individuals it is possible to assume the value of the eligibility variable. If a young man participates in the unemployment assistance program but not the unemployment insurance program, we can assume that he is ineligible for insurance payments. We can make this

assumption because assistance payments are always less than insurance payments (except in the U.K. where the two programs have equal benefits) and the stigma of the assistance program is greater than that of the insurance program, as the benefits of the latter program are seen to be earned (Heintze, 1999). This implies that an individual is highly likely to choose insurance benefits over assistance benefits. On the other hand, if a young man took insurance payments in the past year then we will assume that he is eligible for such payments. Thus the young men for whom we are unable to assess their insurance eligibility status are those who received neither insurance nor assistance payments (that is, 80% of sample). The eligibility dummy is imputed for these individuals.

The technique used is multiple imputation as first developed by Rubin (1987). By this method each missing observation is replaced with K values drawn from a distribution of possible values (Freedman, 1990). In this analysis, five values are drawn for each missing observation (that is, K = 5). Thus five complete datasets are formed, containing five (possibly) different values for the variable 'insurance eligibility' for those individuals for whom the variable is unknown and the same (known) value for eligibility for those for whom the variable is known. The models are run with each dataset and the resultant coefficients (β_k, k = 1 .. K) are averaged to produce one set of results for each model (β^*). The standard errors for these coefficients (SE^*) are calculated using the following formula (Rubin, 1987):

$$SE^{*2} = \text{var}(\beta^*) = \{ \frac{1}{K} \sum_k S_k \} + \{ \frac{K+1}{K} [\sum_k (\beta_k - \beta^*)^2 / (K-1)] \} \qquad \text{(I-4)}$$

where β_k are the coefficients obtained from using the kth dataset (k = 1 .. K), S_k is the variance for coefficients β_k, β^* is the mean of the K coefficients β_k, and K is the number of draws or datasets (in this case 5). Note that the first term of equation I-4 is the mean of the K variances (the 'within sample' variance) and the second term inflates the variance to reflect the uncertainty that

occurs due to the imputation process. It is the 'between' variance enlarged by the term (K+1)/K (Freedman & Wolf, 1995, Wolf, 2000).

How are the unknown values for 'insurance eligibility' derived? The variable that has missing data (in this case, E_I) is assumed to be related to a vector of observed variables T:

$$E_I = f(\theta^*, T, \mu) \qquad \text{(I-5)}$$

where θ^* is a vector of unknowns and μ is a random error term. In our case equation I-5 is determined by a probit model[11]. The probit model implies that we assume that μ is normally distributed. Equation I-5 is used to predict the unknown values of E_I.

Multiple imputation, however, allows for both deterministic and stochastic elements in the estimates of unknown E_I. Note that there are two sources of randomness in generating each imputation of E_I. The first arises from the sampling distribution of θ^* and the second from the random variation in μ (Freedman, 1990). In the case of the first source of randomness, the estimates θ^* are distributed normally with a mean of θ (the true value) and variance Σ. We thus wish to estimate $\hat{\theta}$ (an estimate that includes the randomness of the sampling distribution of θ) from the multivariate normal distribution (θ^*, Σ^*)[12]. If A is the Cholesky root of Σ^*, that is $AA' = \Sigma^*$, and Z is a vector of random numbers normally distributed (that is, $Z \sim N(0,1)$), then:

$$\hat{\theta} = \theta^* + AZ \qquad \text{(I-6)}$$

is distributed normally with mean θ^* and variance Σ^*, since:

$$\text{var}(\hat{\theta}) = A\ \text{var}(Z)\ A' = A\ I\ A' = A\ A' = \Sigma^*$$

[11] Although the eligibility variable is not missing completely at random (MCAR) it does fulfill the conditions for being missing at random (MAR). That is, the missing values for eligibility do not depend on the observed values (Rubin, 1987; Schafer, 1997).

[12] θ^* and Σ^* are unbiased estimates of θ and Σ.

where I is the identity matrix (Freedman, 1990)[13]. Thus for each of the K rounds of imputation the coefficients estimated from the probit are adjusted using (I-6). The probit in question is that whose dependent variable is the known insurance eligibilities.

The second source of randomness arises from the error μ. The probit prediction model allows the dependent variable to equal one if $\theta^{*}T > \mu$ and zero otherwise. Given our estimate $\hat{\theta}$ and the known values for T, the missing values for E_I can be predicted using:

$$E_I = 1 \text{ if } \Pr(\hat{\theta}T) > \mu$$

$$E_I = 0 \text{ otherwise} \qquad \text{(I-7)}$$

where μ is drawn from the uniform [0,1] distribution (Freedman & Wolf, 1995).

To recap, multiple imputation involves five steps. First, the a prediction model for insurance eligibility is formulated, in this case a probit (equation I-5). Second, the coefficients from the probit equation are adjusted to account for the underlying distribution of the coefficients (equation I-6). Third, the imputed value of eligibility insurance is formulated using the adjusted coefficients and the stochastic element of the probit (equation I-7). Fourth, K rounds of imputation are performed, here K = 5, to form K datasets each of which contains the known insurance eligibility and possibly different imputed values for insurance eligibility for those individuals for whom the variable is unknown. Finally, the model is run on all K datasets and the K alternative resultant estimates are averaged with variances that reflect both their within and between sample variance (equation I-4).

One final point to note in this analysis is that one further variable has imputed values. This is a variable that reflects whether or not unemployed young men are currently in school. This is an

[13] Note that the mean of $\hat{\theta}$ must be equal to θ^{*} as the mean of Z (and thus AZ) is zero.

important factor for these men because their age (18-25) makes it likely that some will still be pursuing educational activities. If they are still in school, this must reduce their likelihood of collecting unemployment compensation. This important variable is not available for all individuals but only for 36% (unweighted) of the sample, those in the Canadian subsample.

(ii) Calculation of Trivariate Normals

When using the LIS dataset only three statistical packages are available, SAS, SPSS and Stata. None of these packages contain the software to perform a trivariate probit. It is therefore necessary to calculate a trivariate normal and use maximum likelihood methods to assess the model. Thus it is needful to be able to evaluate the trivariate normal:

$$F_3(x, y, z, \rho_{12}, \rho_{13}, \rho_{23}) = \int_{-\infty}^{h} \int_{-\infty}^{k} \int_{-\infty}^{m} (2\pi)^{-3/2} (\Delta)^{-1/2} \exp\{-\frac{1}{2} [\frac{(1 - \rho_{23}^2)}{\Delta} x^2 + \frac{(1 - \rho_{13}^2)}{\Delta} y^2$$

$$+ \frac{(1 - \rho_{12}^2)}{\Delta} z^2 + 2\frac{(\rho_{13}\rho_{23} - \rho_{12})}{\Delta} xy + 2\frac{(\rho_{12}\rho_{23} - \rho_{13})}{\Delta} xz + 2\frac{(\rho_{12}\rho_{13} - \rho_{23})}{\Delta} yz] \} \, dx \, dy \, dz$$

where $\Delta = 1 - \rho_{12}^2 - \rho_{13}^2 - \rho_{23}^2 + 2\,\rho_{12}\,\rho_{13}\,\rho_{23}$ (Steck, 1958)

I chose to evaluate the trivariate normal using gaussian quadrature methods, specifically the hermite formula. In the simplest terms, gaussian quadrature is the approximation of an integral by a polynomial:

$$\int_a^b w(x)\, f(x)\, dx = \sum_i A_i\, f(x_i) \qquad i = 1, .. n$$

Gaussian quadrature differs from other numerical integration methods in that it chooses the n points for evaluation in an optimal, rather than equally spaced, way. In other words, the nodes x_1 x_2 .. x_n and the weights A_1 A_2 ... A_n are chosen to minimize the expected error in the above approximation. The specific form of gaussian quadrature used in this instance, the hermite formula, is of the form:

$$_{-\infty}\int^{\infty} \exp\{-x^2\}\ f(x)\,dx = \sum_i A_i\, f(x_i) \qquad i = 1, .. n$$

where

$$A_i = \frac{2^{n-1}\, n!}{n^2\, [H_{n-1}\, t_i\,]^2}$$

and t_i is the i-th zero of the hermite polynomial $H_n(t)$. Tables of the nodes and weights from n = 2 to n = 136 are readily available in Stroud and Secrest (1966).

The hermite formula does not exactly equate to the normal distribution. The limits of integration are different, $(-\infty,\infty)$ as opposed to $(-\infty, h)$ in the univariate case[14]. Additionally, we need to transform the function in order to present it in the form: $\exp\{-x^2\}\, f(x)$ in the univariate case[15]. Furthermore, we need to expand the formula to assess a triple, rather than single, integral.

Considering first the univariate normal, that is a single integral, a suitable transformation of the limits of integration on the hermite formula has been performed by Steen, Byrne and Gelbard (1969). They present nodes and weights from n = 2 to n = 15 for the function:

$$_0\int^{\infty} \exp\{-x^2\}\ c(x)\,dx = \sum_i A_i\, c(x_i) \qquad i = 1, .. n$$

Further nodes and weights (up to n = 35) are offered by Yang (1998).

Again considering the univariate case, we next need to transform the hermite weights to assess the univariate formula:

$$f(x) = (2\pi)^{-1/2} \exp\{-1/2\,[x^2]\}$$

If h(x) is a suitably regular function then:

[14] In the trivariate case the limits of integration are $(-\infty,\infty)$, $(-\infty,\infty)$, $(-\infty,\infty)$ as opposed to $(-\infty, h)$, $(-\infty, k)$, $(-\infty, m)$.

$$g(x) = h(x)\,(2\pi)^{-1/2} \exp\{-1/2\,[x^2]\}$$

and:

$$_0\int^{\infty} g(x) = {}_0\int^{\infty} (\pi)^{-1/2}\, h([2]^{1/2} x)\,(\exp\{-x^2\})$$

Using the hermite formula, we obtain:

$$_0\int^{\infty} g(x) = \sum_i (\pi)^{-1/2} A_i\, h([2]^{1/2} x) \qquad i = 1, \ldots n$$

$$= (\pi)^{-1/2} \sum_i [2]^{1/2} (\exp\{-x^2\}) A_i\, g([2]^{1/2} x) \qquad i = 1, \ldots n$$

$$= (\pi)^{-1/2} \sum_i m_i\, g(t_i) \qquad i = 1, \ldots n$$

where $m_i = [2]^{1/2} (\exp\{-x^2\}) A_i$ and $t_i = [2]^{1/2} x$ (Naylor and Smith, 1982)

Expanding this to multiple integrals, in our case to triple integrals, as the trivariate normal forms a three-dimensional simplex, the quadrature formula for the three-dimensional case can be formed from the product of three one-dimensional formulae. That is, we can form the product of three one-dimensional formulae:

$$_a\int^{b} w(x)\, f(x)\, dx = \sum_i A_i\, f(x_i) \qquad i = 1, .. n$$

to produce:

$$_a\int^{b}\,{}_a\int^{b}\,{}_a\int^{b} w(x)\, w(y)\, w(z)\, f(x, y, z)\, dx = \sum_p B_p\, f(x_i, y_j, z_l)$$

$$B_p = A_i\, A_j\, A_k \qquad p = 1, \ldots. n^3$$

(Stroud & Secrest, 1969).

Thus our trivariate normal can be approximated by the polynomial:

$$F_3(x,y,z, \rho_{12}, \rho_{13}, \rho_{23}) = (\pi)^{-3/2}\, \Delta \sum_p m_i\, m_j\, m_l\; g(r_i, s_j, t_l)$$

where $m_i = [2]^{1/2} (\exp\{-x^2\}) A_i$

$m_j = [2]^{1/2} (\exp\{-y^2\}) A_j$

$m_l = [2]^{1/2} (\exp\{-z^2\}) A_l$

---

[15] or $\exp\{-x^2 - y^2 - z^2\}\, f(x,y,z)$ in the trivariate case.

and $r_i = [2]^{1/2}\, x$

$s_j = [2]^{1/2}\, y$

$t_l = [2]^{1/2}\, z \qquad p = 1, \ldots n^3 \qquad (I\text{-}8)$

The last task to evaluate the trivariate normal is to find the appropriate form g (r_i, s_j, t_l) of the trivariate normal polynomial, and transforming the integration such that it is over the interval (0, ∞). Now:

$$F_3(x, y, z, \rho_{12}, \rho_{13}, \rho_{23}) = \int_{-\infty}^{h} \int_{-\infty}^{k} \int_{-\infty}^{m} (2\pi)^{-3/2} (\Delta)^{-1/2} \exp\{-\tfrac{1}{2} [\tfrac{(1 - \rho_{23}^2)}{\Delta} x^2 + \tfrac{(1 - \rho_{13}^2)}{\Delta} y^2$$

$$+ \tfrac{(1 - \rho_{12}^2)}{\Delta} z^2 + 2\tfrac{(\rho_{13}\rho_{23} - \rho_{12})}{\Delta} xy + 2\tfrac{(\rho_{12}\rho_{23} - \rho_{13})}{\Delta} xz + 2\tfrac{(\rho_{12}\rho_{13} - \rho_{23})}{\Delta} yz] \}\, dx\, dy\, dz$$

where $\Delta = 1 - \rho_{12}^2 - \rho_{13}^2 - \rho_{23}^2 + 2\,\rho_{12}\,\rho_{13}\,\rho_{23}$

I chose the transformation:

$$u = \frac{h - x}{(2\Delta)^{1/2}} \qquad v = \frac{k - y}{(2\Delta)^{1/2}} \qquad w = \frac{m - z}{(2\Delta)^{1/2}}$$

that is:

$x = h - (2\Delta)^{1/2} u, \qquad dx = -(2\Delta)^{1/2} du$

$y = k - (2\Delta)^{1/2} v, \qquad dy = -(2\Delta)^{1/2} dv$

$z = m - (2\Delta)^{1/2} w, \qquad dz = -(2\Delta)^{1/2} dw$

Letting

$$h_1 = \frac{h}{(2\Delta)^{1/2}} \qquad k_1 = \frac{k}{(2\Delta)^{1/2}} \qquad m_1 = \frac{m}{(2\Delta)^{1/2}}$$

Then:

$$F_3(x, y, z, \rho_{12}, \rho_{13}, \rho_{23}) = \Delta\,(\pi)^{-3/2} \int_0^{\infty} \int_0^{\infty} \int_0^{\infty} \exp\{-u^2 - v^2 - w^2$$

$$+ h_1(2u - h_1) - \rho_{23}^2 (h_1 - u)^2 + k_1(2v - k_1) - \rho_{13}^2 (k_1 - v)^2$$

$$+ m_1(2w - m_1) - \rho_{12}^2 (m_1 - w)^2 + 2 (\rho_{12} - \rho_{13} \rho_{23}) (u - h_1) (v - k_1)$$

$$+ 2 (\rho_{13} - \rho_{12} \rho_{23}) (u - h_1) (w - m_1)$$

$$+ 2 (\rho_{23} - \rho_{12} \rho_{13}) (v - k_1) (w - m_1)\} \, du \, dv \, dw \qquad \text{(I-9)}$$

The trivariate normal is thus in a form that can be evaluated by the quadrature formula outlined in (I-8), where g (u, v, w) is the function appearing after the triple integral in (I-9). (Note that that transformations $r_i = [2]^{1/2} u$, $s_j = [2]^{1/2} v$ and $t_l = [2]^{1/2} w$ need to occur.) The calculation of the trivariate is most accurate when the values of h, k and m are low, as they increase the evaluation becomes less exact. When n=10 nodes are used, the calculation is exact to 4 decimal places when h, k and m are negative, but only to 2 decimal places when all three are large (that is greater than +1). Increased accuracy may be obtained by increasing the number of nodes.

Note that when eligibility for insurance is zero, that is, the individual is not eligible for insurance, the probit to be evaluated is a bivariate probit. We can use the gaussian quadrature formula (I-8) albeit for two integrals, that is:

$$F_2(x,y, \rho_{12}) = (\pi)^{-1/2} (1 - \rho_{12})^{1/2} \sum_p m_i \, m_j \; g (s_j, t_l)$$

where $m_i = [2]^{1/2} (\exp\{-x^2\}) A_i$

$m_j = [2]^{1/2} (\exp\{-y^2\}) A_i$

and $s_i = [2]^{1/2} x$

$t_j = [2]^{1/2} y \qquad p = 1, ... n^2 \qquad \text{(I-10)}$

and the appropriate $g (u, v) = \exp \{-u^2 - v^2 + h_1(2u - h_1) + k_1(2v - k_1) + 2\rho_{12} (u - h_1) (v - k_1)\}$ (Drezner, 1978).

The maximum likelihood required to evaluate the trivariate probit is performed using the ml module in Stata. This module uses a modified Marquardt (1963) algorithm (which in itself is a modified Newton-Raphson method). During each iteration a numerical derivative of the log likelihood function is calculated, and a step of length $H^{-1}g$ is taken. Here H is the negative of the Hessian and g is the gradient. After the step the likelihood is recalculated. If it has increased another step is taken, if decreased a step is taken in the direction of g (Stata Press, 1997).

If we let: $q_{1i} = 2(Y_{Hi}) - 1$

$q_{2i} = 2(P_{Ai}) - 1$

$q_{3i} = 2(P_{Ii}) - 1$

then the log likelihood function takes the form:

$$L = \sum_{i\,(E=1)} \ln F_3 \,(q_{1i}\beta_H X_i\,,\, q_{2i}\beta_A X_i,\, q_{3i}\beta_I X_i,\, q_{1i}q_{2i}\,\rho_{12},\, q_{1i}q_{3i}\,\rho_{13},\, q_{1i}q_{2i}\,\rho_{23})$$

$$+ \sum_{i\,(E=0)} \ln F_2 \,(q_{1i}\beta_H X_i\,,\, q_{2i}\beta_A X_i,\, q_{1i}q_{2i}\,\rho_{12})$$

where E is insurance eligibility.

(iii) Predicted Wage and Standard Error Recalculation

One of the dependent variables in the trivariate probit we wish to evaluate is the incidence of work. An important variable in this estimation is the individual's wage. We can theorize that if the individual has the potential to earn a higher wage he is more likely to work as his alternative sources of income are less remunerative in relative terms. However, we do not observe the wage for those who are not working and must therefore predict it. We achieve this with a standard model where the hourly wage is regressed on a set of independent variables for those for whom the wage is known. The set of independent variables includes an inverse mills ratio calculated from a previous equation that controls for whether the individual is working. The wage

regression is then used to predicted values of the hourly wage for the entire sample (Killingsworth, 1983).

When we include the predicted wage in the trivariate probit we are including in this estimation a variable that is measured with sampling error thus biasing the covariance matrix of the trivariate probit. We thus need to correct the standard errors of the trivariate probit estimation. This will be achieved using an extension of the method outlined in Murphy and Topel (1985). Before we correct the standard errors in the third and final stage of the analysis (the trivariate probit) we first need to correct them in the wage prediction regression (the second stage) as this, too, contains a variable from a previous stage (the first stage), that is, the inverse mills ratio.

Note that the partially imputed variables 'insurance eligibility' and 'in school' are included in the trivariate probit. The standard errors are, however, not corrected for these variables. They are not corrected in the final stage because the missing values of these variables are imputed and the known values are included in the data rather than all elements being estimated. The technique of multiple imputation implies that the missing values are not being estimated but rather are simulated and thus the standard errors do not require re-evaluation as multiple imputation allows for the sampling error included in the coefficients of the estimating equation. Why then are the "missing" wage values for non-workers not imputed rather than estimated? They are not imputed because the unknown values of the wage cannot be measured, they must be estimated. Multiple imputation is used to generate missing values for variables that can be measured, but for some reason were not[16]. As the probits used in the imputation of 'insurance eligibility' and 'in school' are not used in estimation, the standard errors remain uncorrected.

[16] Thanks to Doug Wolf for clarification of this insight.

The regression used to predict the wage contains a variable from a proceeding probit. Thus equation (24) in Murphy and Topel (1985) is used to correct the errors in this second stage. In the final stage however, the estimated variable is forwarded from a regression into a maximum likelihood estimation and Murphy and Topel do not include such a format in their study[17].

In the case of two regression analyses they estimate the distribution of the estimators in the latter stage by:

$$(n)^{1/2} (\theta_b^* - \theta_b) = (n^{-1} Z'Z) (n^{-1} Z'F^*) (n^{-1} (\theta_a^* - \theta_a)) + (n^{-1} Z'Z) (n^{-1/2} Z'U) \qquad \text{(I-11)}$$

where θ_a are the coefficients in the former stage and θ_b in the latter stage, the true values are indicated by the asterisk, Z is the matrix of independent variables in the latter stage (including the estimated values from the previous stage) and the elements of F^* are:

$$f_{ij} = \gamma (\partial f / \partial \theta_{bj})$$

where γ is the coefficient of the estimated variable in the latter stage, f is the function used to predict the variable being forwarded to the latter stage, and the partial derivative is evaluated for each of the j independent variables in the former stage. (Note that here we assuming that only one estimated value is included in the latter stage.) Murphy and Topel show that $(n)^{1/2} (\theta_b^* - \theta_b)$ is asymptotically normal with mean vector zero and covariance matrix:

$$V_b = \sigma^2 Q_0^{-1} + Q_0^{-1} Q_1 V_a Q_1 Q_0^{-1} \qquad \text{(I-12)}$$

where:

$$Q_0 = n^{-1} Z'Z$$

$$Q_1 = n^{-1} Z'F^*$$

and V_a is the covariance matrix from the former stage.

[17] In the following discussion, the first stage is the initial probit from which comes the mills ratio, the second stage is the regression, which calculates the predicted wage, and the third stage is the maximum

When estimating a maximum likelihood we have (in general) the following asymptotic equivalence:

$$(n)^{1/2} (\theta^{*} - \theta) = - R^{-1} \sum_{i} n^{-1/2} (\partial L / \partial \theta)$$

where θ is the vector of coefficients estimated, L is the likelihood and R is Fisher's information matrix:

$$R = - E[n^{-1} (\partial^2 L / \partial \theta \, \partial \theta \text{ '})$$

The equivalent equation of I-11 for forwarding a least squares estimation into a maximum likelihood (stage two into stage three) is thus:

$$(n)^{1/2} (\theta_3^{*} - \theta_3) = R_3^{-1} (\sum_{i} n^{-1/2} (\partial L_3 / \partial \theta_2)' F_2^{*}) (n^{-1} (\theta_2^{*} - \theta_2)) - R_3^{-1} \sum_{i} n^{-1/2} (\partial L_3 / \partial \theta_3)$$

where R_3 is Fisher's information matrix for the third stage and L_3 is the likelihood for the third stage. The asymptotic distribution of the third stage estimators is normal:

$$(n)^{1/2} (\theta_3 - \theta_3^{*}) \sim N(0, V_3)$$

$$V_3 = R_3^{-1} + R_3^{-1} Q_2 V_2 Q_2 R_3^{-1}$$

where $Q_2 = n^{-1} (\partial L_3 / \partial \theta_2)' F_2^{*}$ and V_2 is the covariance matrix calculated at the second stage using Murphy and Topel's equation (24).

I-VI: Results

(i) Descriptive Results

Descriptive statistics exploring the men's behavior in the labor market are shown in table I-4. Older men (26-65) are more likely to be active in the labor market than young men in all three countries are. Young men are most likely to be in the labor market in the U.K. (83 percent) and least likely in Canada (72 percent). Once in the labor force and thus either employed or actively looking for a job, young men are more likely to be unemployed than are older men. In each

---

likelihood estimation of the trivariate probit. Stages a and b are theoretical first and second, former and latter, stages.

country young men are unemployed at over twice the rate of their older counterparts. American youth have the lowest unemployment rate at 11 percent and Canadian the highest at 17 percent.

<Table I-4 about here>

The next statistics measure the rate of unemployment compensation (UC) receipt. The first three statistics of the set indicate benefit receipt in the past year. Upon naïve examination of the first three lines of this second section of table I-4, it appears as if, in Britain and Canada, young men are far more likely to be compensated for their unemployment than are older men. Young men are more likely than older men to receive assistance or insurance benefits in these two countries. However, as noted above, young men are far more likely to be unemployed. A more useful statistic is thus current receipt among those currently unemployed. Due to data restrictions, we have no measure of current receipt. The last line in this second section of table I-2 measures past unemployment compensation receipt for those unemployed at the present time. The singular nature of this measure produces difficulties in comparing results across countries, difficulties further discussed in the following section. Given this caveat, and interpreting this variable naively as current unemployment compensation receipt among the currently unemployed, it appears that young men are less likely than older men to be compensated for their unemployment in either Canada or America. Only in Britain are unemployed young men more likely to receive benefits than are older men. Again only in Britain do young men have an excellent chance of receiving some form of unemployment compensation. By contrast, young men in the U.S. have the lowest chance of receiving benefits, only one-third of young men are compensated for their unemployment.

Examining unemployment benefits (the sum of unemployment insurance and assistance payments) it appears as if American benefits are far lower than those offered in Britain and

Canada.[18] However, as indicated, Americans are unemployed for fewer weeks. When we consider the average weekly benefit, it can be seen that benefit amounts are less disparate across the nations included. Benefits are most generous in Canada and least generous in the U.K.. Older men receive more per week than do younger men in each country. This most likely occurs because many unemployment assistance and insurance programs offer extra benefits for dependents and older men are likely to have more dependents. Additionally, older men tend to have higher earnings and most unemployment insurance programs pay higher benefits to higher earners. Young men, if they experience unemployment at all, are on average unemployed for the least time in the U.S. (20 weeks) and the most time in the U.K. (32 weeks).

The last set of measures in table I-4 concern labor earnings and income. Median wage and income amounts are offered, rather than means, due to outliers at the upper end of the distribution. The median wage for working young men is lower than that of working older men in each of the three countries. The median for Canadian male youth is the highest ($9.48) and British youth have the lowest ($6.64). This ordering is reversed when we consider the average hours worked per week. As with median wages, median income per household is higher for those households that contain older men than for those that contain younger men. British young men live in households with lower incomes than their Canadian and American counterparts and Canadians youth are attached to households with the highest median. A striking result is found when we examine the percent of household income that is provided by the unemployment compensation benefit. On average, between 3.5 and 5 percent of British and Canadian household income is provided by unemployment benefits; the equivalent American figure is 1.5 and 2.25 percent. This indicates that American male youth do not rely as much on unemployment benefits for support as do youth in the other two countries.

[18] All monetary amounts have been converted into 1995 American dollars using purchasing power parities.

(ii) Multiple Imputation Results

Table I-5 contains the probit equation used to give the initial values of insurance eligibility for those for whom this is unknown. It is estimated on those for whom insurance eligibility is known (unweighted N = 1979). Although none of the programmatic variables included are statistically significant, the dummy variables for nations (Britain and Canada, with the U.S. as the excluded category) indicate national differences in the probability of being eligible for unemployment insurance. We infer from the descriptive results that British young men are less likely to receive unemployment insurance than are American young men, while Canadian young men are the most likely to receive insurance of the three nations studied. It would be easy to interpret this as being due to programmatic differences as the British insurance system has a long qualification period for eligibility and the Canadian system the shortest, yet these programmatic differences are included in the analysis in table I-5 but the variables are not significant. The national differences must therefore be due to other causes – perhaps differences on a bureaucratic or administrative level. Bureaucratic barriers are theorized to have an effect on receipt of welfare benefits (Meyers, Heintze and Wolf, 2002)[19].

< Table I-5 about here >

Eligibility for insurance is also affected by the individual's self employment status, with those self employed being less likely to be eligible for benefits. As insurance contributions (and thus eligibility for receipt of benefits) rely on both the employee and employer, it unsurprising that those who are self employed are less likely to be eligible. While the system allows for the

[19] Although we are considering eligibility for benefits rather than receipt of benefits, our definition of eligibility is predicated on receipt.

individual to undertake both portions of the insurance contributions, the self-employed may chose to declare less income, or actually earn less income[20], which may affect eligibility.

A series of occupational dummies are included in the analysis to test whether eligibility is affected by the individual's type of employment. Occupation may affect eligibility in that some occupations have longer and steadier employment than others, and involuntary layoffs affect insurance eligibility. The only occupation or industry that appears to effect eligibility is the service industry. This may be a reflection of the age of the sample. It may be that these young men are more likely to have experienced longer term employment in the service industry by holding a casual job at a younger age, or while in school in the past.

The estimation used to impute whether the individual is in school or not is included in table I-6. Many of the demographic variables are statistically significant and display the expected signs. The variables that indicate that the individual has embarked on a more settled, longer-term life style display a negative relationship with the dependent variable. In other words, those who are married or co-habitant, those who are older and those who have left their family of origin are less likely to be in school at the current time. In the same way, those who have never held a job, or who have not held one for the past 5 years, are more likely to be still in school.

< Table I-6 about here >

One less obvious result is the affect of the number of earners in the family. The fewer earners there are, the more likely it is that the individual is in school. One explanation may be that there

[20] Contradictory reports exist as to whether the self employed earn more or less than traditional wage and salary employees, with one author suggesting that the differences are based on the choice of the functional form for the earnings equation (Portes & Zhou, 1996). However, a time series analysis of earnings of the

is some kind of need or modeling effect. Households where there are more earners encourage the young men to work rather than to attend school. Alternatively, perhaps households with more workers are those households, which require more income or have higher levels of need. It may be that the work effort of the young man is needed for the income he can supply rather than encouraging him to attend school. Note, however, that this latter explanation becomes less feasible when we realize that the level of other income in the household, as divided by an equivalency scale has no significant affect on the dependent variable. The last significant variable is that of rural residence. Those who live in less densely settled areas are less likely to be in school.

Imputation of unknown values for insurance eligibility and being in school are performed five times. As indicated in the econometric specification section above, the each imputation contains stochastic and deterministic elements. Table I-7 shows the means and standards of both variables in each imputation run. Over fifty percent of the sample is eligible for insurance once imputation is undertaken. This is a higher percentage than occurs in the subsample where eligibility is known, but as those who have known eligibility are less likely to be working, this higher rate of eligibility is not surprising as eligibility is obtained through longer work history.

< Table I-7 about here >

The results from the imputation of those in school are also included in table I-7. After imputation, between four and four and one-half percent of the sample is in school[21]. Again this is

British self-employed indicates that income among the self-employed has been on the decline since 1980 in that country (Robson, 1997).

[21] This may appear to be extremely low given the age range of the sample, but we must note that the sample includes only those in the labor market and we are measuring the school status of those who are unemployed. Those going to school full-time who neither hold, nor are looking for a job, are not included in the analysis.

a higher percentage than occurs in the subsample where school attendance is known, but the subsample (Canadians) are less likely to complete tertiary education than are the British or American young men. School attendance is not an available measure in the LIS dataset for Americans, but it is available in the CPS, the dataset from which the American LIS data is drawn. When a subsample of men aged between 18 and 24 is drawn from the 1994 CPS sample[22], the equivalent in school percentage is 0.029, very similar to both the Canadian measure and the imputed full analysis sample measure.

(iii) Wage Prediction

The prediction equation for the hourly wage is displayed in table I-8, which includes standard errors corrected for the presence of the inverse mills ratio from an earlier probit (not displayed)[23]. This inverse mills ratio controls for the estimation being performed only on workers. Many of the demographic variables display statistical significance and the effects are in the expected direction. Those with less education (education levels 1 and 2, the best equivalent to less than a high school education and a high school diploma) have lower wages than those in the omitted category, that is those with a tertiary degree or higher. Interestingly, those who have started, but not completed, tertiary studies have wages that are not significantly different from those with a degree. Older individuals have higher wages than younger ones perhaps indicating greater work experience. Note that we have not included the square of age in the analysis, usually included to model the curvature of the effect of age on wage, because the sample is very age limited. All included individuals are between the ages of 18 and 25.

< Table I-8 about here >

[22] Unfortunately, school attendance is not available for those aged 25.
[23] Note that the correction to the covariance matrix must be performed as the covariance matrix from the wage prediction estimation is forwarded to the standard error correction performed on the covariance matrix of the trivariate probit in the final stage.

When we examine the national and regional variables, we see that British young men have significantly lower wages than Americans (the omitted category) do, while there is no significant difference between the wages of Canadians and Americans. Those who live in rural areas have significantly lower wages than more urban dwellers. Those who live in areas of higher unemployment have lower wages reflecting the greater supply of labor. Included in the estimate is a measure of the regional male median wage. Although this coefficient is not statistically significant it is used to identify the instrumented wage that is used in the next stage of the model. It could be assumed that the coefficient on male median wage is biased because of the weak correlation between the male median wage and the individual wage (Bound, Jaeger & Baker, 1995), however Angrist and Krueger (1999) have shown that the "weak instruments" problem is not of concern for just identified two-stage estimations. Furthermore, bias is further guarded against in that for 14 percent of the sample the instrumented is estimated on a non-intersecting sample.

(iv) Joint Estimation of Work, Unemployment Assistance Receipt and Unemployment Insurance Receipt

The results for model 1 are outlined in table I-9. Our discussion will focus on the variables of interest, the programmatic variables, only those demographic and regional variables that attain statistical significance will be discussed. No demographic variable that attains statistical significance has an unexpected sign, excepting the impact of tax on work. Those who are currently in school are less likely to work, an unsurprising outcome given the time constraints that school attendance imposes. Those with a education level of 1 (which may be considered as having the equivalent of less that a high school education) are less likely to work that those of education level 4 (completed tertiary studies). Young men of lower education are less likely to be attractive to prospective employers. Young men who have embarked on a more settled lifestyle,

that is they are married and head their own households, have a higher probability of working. Those who are self-employed are more likely to be working, again an unsurprising result as few individuals who consider themselves to be self-employed will describe themselves as not working. There appears to be an affect of modeling in employment – those who live in households with more workers are more likely to be working themselves. Yet living in a household with more equivalent household income lessens the probability of work for young men. Two regional variables attain statistical significance – young men who live in areas with higher unemployment are less likely to work. More surprisingly, those in regions with high tax rates are more likely to work.

The impacts of the demographic and regional variables on the two welfare programs are as may be expected. Those in school have a lower probability of receiving assistance benefits, perhaps because most individuals who attend school plan for the costs they must face. Young men in households with more children are both more likely to receive assistance and more likely to receive insurance. This may be because the needs of the children encourage those eligible for benefits to apply for them. Unsurprisingly, a lower number of earners in the household also encourages use of both assistance and insurance programs. Stability of housing encourages young men to apply for assistance, a result found elsewhere (Berger et. al., 2001). Young men who are self employed are also more likely to receive assistance benefits, perhaps because this work is harder for the authorities to monitor. Young men who head their own household have a higher probability of receiving insurance benefits. Finally, lower tax rates encourage young men to receive insurance benefits.

<Table I-9 about here>

With regard to the programmatic variables, let us first consider the affects on the decision to work. A higher maximum assistance benefit reduces the probability that the individual will work. On the other hand, insurance benefits do not have the same effect. The maximum insurance benefit has no affect on the individual's decision to work. This indicates that while assistance benefits have an expected disincentive effect on work, the same is not true of insurance benefits. The last programmatic variable, the maximum duration of UI effects, has the largest effect on the decision to work, but is only significant at the eleven percent level in a two-tailed test. However, the sign of the coefficient indicates that the longer it is possible for an individual to collect benefits, the more likely he is to be working at the current time, given his decisions about collection of unemployment compensation in the previous year

The effects of the programmatic variables on assistance receipt are as expected. The higher the maximum assistance benefit, the more likely the individual is to choose to collect assistance payments. On the other hand, the shorter the time one is allowed to collect insurance benefits, the more likely the individual is to choose to collect assistance. As the size of insurance benefits decrease, young men are more likely to receive assistance.

Finally, we examine the effects of programmatic variables on insurance receipt. The only significant variable is that of the maximum assistance benefit. Given that the individual is eligible for insurance benefits, the higher the maximum assistance payment, the more likely the individual is to receive insurance benefits. This appears to be inexplicable but most likely relates to the issue of stigma. It appears that when one is eligible for the less stigmatized insurance benefits, the possibility of increased assistance benefits does not discourage the receipt of insurance benefits, even while controlling for the maximum amount of the insurance benefits. The maximum amount of the insurance benefit itself, does not affect the receipt of insurance. It

is also true that, given eligibility for insurance, work decisions and assistance receipt decisions, insurance receipt is unaffected by the maximum length of time that benefits can be received.

Note that all three of the correlations in table 9 attain statistical significance. This implies that controlling for the relationship between the errors in the estimation was important, justifying the use of the trivariate probit.

The second model, the results of which are displayed in table I-10, introduces further programmatic variables - the amount of the insurance and assistance benefit that can be retained before the benefit is reduced due to other income. We will limit our discussion to two areas: one, variables whose statistical significance is different from the first stage; two, the programmatic variables.

< Table I-10 about here >

Comparing the results of models one and two among the demographic and regional variables, the results are similar. The major differences lie in the coefficients for the demographic variables describing the receipt of insurance benefits. In model two, the predicted wage and education have an statistically significant impact on insurance receipt. Those with higher wages are more likely to receive insurance benefits, as are those of lower education. This latter may seem counter-intuitive, but given that we are controlling for the wage, it may be that those of lower education are more likely to apply for benefits than those of higher education – and are thus more likely to receive benefits.

Although introduction of the additional programmatic variables has little effect on the demographic and regional variables, the effects of the programmatic variables are affected. As

with the previous model, an increase in the maximum assistance benefit reduces the probability that the individual will work, but the effect is no longer significant at the five percent level. On the other hand, increasing the maximum insurance benefit now appears to reduce the chance that a young man will work indicating unemployment insurance benefits act as a work disincentive. However, alteration, and indeed liberalization, of other aspects of the unemployment insurance program appear to increase the possibility of employment. Increasing the amount of the benefit that a young man may retain (before his earnings require reduction of the benefit) appears to encourage work.

The programmatic variables' effects on assistance receipt are as expected. Given the individual's decisions about work and insurance, increasing the maximum allowable assistance benefit increases the probability that the young man will participate in the assistance program. Although significant only at the eleven percent level, increasing the maximum insurance benefit reduces the chance of assistance program participation.

When we move to consideration of participation in the insurance program, we find that all of the programmatic variables have a significant influence on the dependent variable. As with the previous model, increasing the maximum assistance benefit increases the chance that the individual will participate in the insurance program given that he is eligible to do so. As may be expected both increasing the maximum insurance benefit and increasing the maximum amount of time that benefits can be received also encourages participation in the insurance program (given eligibility). The probability of participation in the insurance program is increased, however, by reducing the amount of the benefit that is unaffected by small increases in other forms of income.

I-VII: Discussion and Conclusion

Our descriptive statistics appear to indicate that American men between the ages of 18 and 25 are very unlikely to receive compensation for unemployment they may experience. This is particularly true when we compare young American men's situations with those of young men from Britain and Canada. According to our metric, only one-third of unemployed American young men can expect to receive benefits while unemployed, compared to forty-five percent of Canadian young men and ninety percent of British young men. However, we must take into account the peculiar nature of the measure used. We measure past receipt for those currently unemployed. If we can assume that those who collected unemployment compensation in the past year also collect it at the current time, then our measure is somewhat straight-forward. However, duration of eligibility for unemployment insurance does vary by country. Britain has the longest possible duration and America the shortest. If individuals collected unemployment insurance in the past year, British young men are more likely to still be collecting it at the current time than are Americans. We may therefore be underestimating the percentage of American young men who are currently compensated for their unemployment as compared with British and Canadian young men. If we can assume that all unemployed young men who lost their eligibility for unemployment insurance benefits through time limitations chose to receive unemployment assistance in its place, then our measure is still accurate (and does not underestimate American receipt compared to British and Canadian receipt). The stigma attached to assistance receipt, however, makes this eventuality unlikely.

Canadian unemployed young men are more likely than Americans to receive benefits through the unemployment insurance program rather than the assistance program. In the United States the chances of receipt are almost identical between the programs. Yet the distinction between them is important, particularly in Canada and the U.S..

Why is participation in insurance programs as opposed to the assistance programs important? Participation in the insurance scheme carries less stigma than the assistance scheme and the benefits are higher[24]. Higher benefits allow young men and their families to meet their needs more easily. As young men, when they are employed, earn less than older men they are less able to cover any disruption in income that occurs while they are not working as they are less likely to hold assets or have any considerable savings[25]. Higher benefits are thus very important to young men. On the other hand, we must remember that unemployment insurance is only paid to those who are unemployed through no fault of their own. Eligibility criteria are therefore harder to fulfill than are those for unemployment assistance programs.

Why are American young men less likely than Canadians to receive benefits through the unemployment insurance scheme? American men are less likely to receive insurance benefits than Canadian men because in the U.S. there exists an earnings requirement over a particular period rather than just a simple work duration requirement. Canadians are vested, if new entrants in the workforce, after 26 weeks of work. Canadian workers with longer employment histories have an even shorter qualification period. All Americans must earn a certain amount, dependent on the state, and generally wages must be earned in at least two quarters to be eligible for insurance. (Although Britain has a longer required work period than either Canada or the U.S. this system is better complemented by a means-tested assistance program, a program that has benefits equal to those of the insurance program.)

Yet this research also offers a more fundamental reason why we should encourage young men to participate in insurance rather than assistance programs given eligibility. Like other populations who receive benefits under means-tested programs (for example, see Moffitt, 1983), assistance

[24] The distinction is less marked in Britain where the benefit levels between the two programs are the same and are administered through the same bureaucratic offices.

benefits clearly act as a work disincentive. Higher benefits reduce the probability of work. However, it is not clear that insurance benefits act in the same manner. Raising the maximum insurance benefit had no statistically significant effect on work participation probabilities in the first model examined here. Under the second model, increasing the maximum insurance benefit does appear to reduce the probability of work, but the effect is much smaller than that of increasing the unemployment assistance benefits.

Our examination of the programmatic elements of the unemployment compensation programs should not be limited to the level of benefits they offer. We should examine the effects of programmatic variables that increase the probability of working and decrease, or leave unchanged, program participation. Two variables fit this description, one in each model. In the first model, increasing the maximum length of time that benefits can be received increases the probability of working, decreases the receipt of assistance benefits, and has no significant effect on the probability of receiving insurance. We theorize that increasing the length of time an individual can receive benefits allows the young man to make a more extensive job search and thus secure a better, longer lasting job. In other words, given decisions about receipt of insurance in the past year if eligible (and decisions about assistance receipt), increasing the time in the past year that the young man has to search for a new job through the support of unemployment insurance encourages employment at the current time. Equally, being entitled to insurance benefits for a longer time makes it less likely that the individual will chose to participate in unemployment assistance programs as to chose assistance benefits over insurance benefits, if eligible, is illogical.

Table I-11 displays the predicted probability of working under different values of maximum benefit duration. All other variables are held at their sample mean. If the maximum benefit

[25] Although it is true that they are likely to have fewer commitments for their income.

receipt duration were at its mean value (31 weeks), the model predicts that 87.1 percent of young men would work. Varying this duration from 20 weeks to 40 weeks implies that the predicted probability of working ranges from 85.1 percent to 88.7 percent or 3.6 percentage points. While this may not seem very large, a four percentage point change in the probability of work implies that approximately 742,500 young men are moving into the workforce. However, since the average weekly insurance benefit is US$263, if every young man stayed on unemployment insurance for the full extra 20 weeks, the 20 week increase in benefit duration (which predicts a four percent increase in employment) could cost approximately US$5,260 per individual. A more realistic approximation of cost may be obtained by taking the average percentage of maximum possible benefit duration that young men actually take. On average young men take 71 percent of the maximum time allowable. Under this metric the approximate cost of increasing the maximum duration by 20 weeks is US$3,682 per individual[26].

<Table I-11 about here >

In the second model, a different variable fulfills our criteria: the insurance benefit disregard. If we increase the amount of the insurance benefit that the individual keeps before the benefit is reduced due to other income then the individual is more likely to work, less likely to receive insurance, if eligible, and it has no affect on assistance receipt. This change in the program makes it more likely that a young man can take a part-time job while still receiving benefits. Enabling him to do this not only increases his income and encourages work behavior, but also improves his work history, making him a more attractive employee in the future. As the insurance benefit is time limited, there is no possibility of the young man continuing to collect benefits and maintaining this part-time job.

[26] Note that this is a very approximate measure that assumes that all other factors remain constant while only the duration that benefits can be received is increased. In particular, this rough estimate assumes that

The second model predicts the probability of work as 87.6 percent at the mean values of the independent variables. Table I-11 indicates the effects of varying the insurance benefit disregard while holding all other variables at the mean. Decreasing the benefit disregard to $40 per week (from its mean value of $58) reduces the probability of work to 87 percent; increasing the benefit disregard to $80 per week increases the percentage of young men that will work to 88.2 percent.

It appears as if increasing the maximum duration that the individual receives insurance has more effect on employment than does increasing the insurance benefit disregard. However, this is merely the effect of the range of values displayed in table I-11 for the variable in question. If we increase the benefit disregard to $170 per week, then the second model predicts that employment will increase by 3.6 percent, holding all other variables at their mean. If young men continue to spend an average of 22 weeks collecting benefits and all young men take advantage of the full $130 increase in the benefit disregard, then the cost of this program change will be approximately $2,860 per individual.

It thus appears as if increasing the benefit disregard is the less expensive, and perhaps the more attractive, alternative for increasing the percentage of young men who are employed. However, the estimations made above are extremely rough and can only be taken as an approximate guide. Note that although an increase of 20 weeks in the duration of benefits falls into the range of data included in the estimation, an increase to $170 in the benefit disregard does not. The maximum benefit disregard in the data set is $166. Furthermore, measurement of the true effects of the benefit disregard is not easy to capture in a single variable. While the variable included in these estimations captures the amount of the benefit that is disregarded before other income causes reduction of the benefit amount, not all insurance schemes reduce on a dollar for dollar basis.

---

young men participate in the insurance program for the same proportion of the maximum time allowable.

The simplified variable included above cannot capture the full effect of earnings (and other income) on the insurance benefit. On the other hand, if we consider increasing the duration of benefits to be a better policy alternative, the variable is not significant in the second model, indicating that, although in a simple model duration changes may be effective, manipulating its value may have less effect in reality.

In summary then, if we wish to manipulate insurance and assistance programs to encourage young men to work and, at the same time, discourage young men from using relying on welfare assistance, then this analysis indicates that the most effective alternatives are to increase the length of time that young men can accept benefits and/or to increase the insurance benefit disregard. However we must remember that this analysis is concentrated on the behavior of young men. There is no indication that these policy alternatives are effective for women or for older men as their situations differ from those of young men. These suggested policy changes liberalize the insurance programs and thus may discourage work behavior for women or for older men. Note, too, that we cannot assume that the effects of unemployment compensation programs found in this study are solely those of American young men.

Why should we be specifically concerned with the work situations of younger men? Young men are just beginning their life in the labor force and, compared to older men, are less likely to be employed. Yet with our aging population it has become increasingly important for young men to work. The lower employment rate of young men indicates that they have more difficulty in finding and retaining work than older men and thus the insurance schemes may have to offer better work incentives for young men. With regard to the situation of young men as compared to women, it may be necessary to make policy changes that encourage young men to work because the work effort of men, despite vast changes in the gender of household workers in the past three

decades, is still vitally important for the well-being of most families. Policy changes that encourage men to start and maintain a steady work history are thus vital.

Can we apply these results on the interaction of unemployment compensation and work specifically to American young men? When I restrict the analysis to American youth and use the resultant likelihood in a likelihood ratio test with the likelihood of the full sample, I reject the hypothesis that the model's parameters will be similar. This is true for both models. In the American-only estimations (not shown) none of the programmatic variables have statistically significant impacts of work participation. However we must be aware that part of being American is to live in the American policy environment. The full sample results, pooled between a sample of 'Anglo' young men, allow us a view, albeit somewhat blurred, of the impact of a policy environment that differs from ours.

It is also worth noting that this analysis examines the effects of unemployment compensation on the work efforts of young men in the labor force, but it says nothing of the effects of unemployment compensation on those who chose not to enter the labor force. If an individual has no legitimate means of support (that is, support through employment or unemployment compensation) he may choose not to enter the labor force, particularly if he expects to experience barriers to employment. Such barriers may include minority status, poor education or skills. If the individual is not in the labor force, he has a far smaller chance of finding employment. Unemployment compensation may play a role in funding job search for such individuals and in encouraging legitimate labor behavior in the future. America's low unemployment compensation coverage for young men compared to Britain and Canada may encourage young men to seek support through less orthodox or less legitimate channels, an idea that will be further discussed in section III of this thesis.

References:

Anderson, P.M. & Meyer, B.D. (1997). "Unemployment Insurance Takeup Rates and the After-Tax Value of Benefits." The Quarterly Journal of Economics, 112 (August), pp.913-937

Angrist, J.D. & Krueger, A.B. (1999). "Empirical Strategies in Labor Economics" in Handbook of Labor Economics, Vol 3A (pp. 1277-1366). Ashenfelter, O.C. & Card, D. (Eds.). Amsterdam, NL: North-Holland

Atkinson, A.B. & Micklewright, J. (1991). "Unemployment Compensation and Labor Market Transitions: A Critical Review". Journal of Economic Literature, 29 (December), pp.1679-1727

Baumol, W.J. (1967). "Macro-Economics of Unbalanced Growth: The Anatomy of Urban Crisis." American Economic Review, 57(3), pp.741-755

Ben-Harim, M. & Zukerman, D. (1987). "The Effect of Unemployment Insurance on Unemployment Duration." Journal of Labor Economics, 5, pp.386-390

Bound, J., Jaeger, D.A. & Baker, R.M. (1995). "Problems with Instrumental Variables Estimation When the Correlation Between the Instruments and the Endogenous Explanatory Variable is Weak." Journal of the American Statistical Association, 90, pp. 443-450

Cage, R. (1994). "How does Rental Assistance Influence Spending Behavior?" Monthly Labor Review, May, pp.17-28

Corak, M. & Jones, S.R.G. (1995). "The Persistence of Unemployment: How Important were

Regional Extended Unemployment Insurance Benefits?" Canadian Journal of Economics, 28(3), pp.555-567

Danziger, S., Haveman, R. & Plotnick, R. (1981). "How Income Transfer Programs Affect Work, Savings and the Income Distribution: A Critical Review." Journal of Economic Literature, 19, pp.975-1028

Drezner, Z. (1978). "Computation of the Bivariate Normal Integral." Mathematics of Computation, 32(141), pp. 277-279

Esping-Anderson, G. (1990). The Three Worlds of Welfare Capitalism. Cambridge: Polity Press

Freedman, V. (1990). "Using SAS to Perform Multiple Imputation." Urban Institute Population Studies Center Working Paper #6. Washington, DC: The Urban Institute

Freedman, V. & Wolf, D. (1995). "A Case Study on the Use of Multiple Imputation." Demography, 32(3), pp. 459-470

Fraker, T. & Moffitt, R. (1988). "The Effect of Food Stamps on Labor Supply: A Bivariate Selection Model." Journal of Public Economics, 35, pp.25-56

Godbout, T.M. (1993). "Employment Change and Sectoral Distribution in 10 Countries, 1970-90." Monthly Labor Review, 113 (October), pp.3-20

Gornick, J.C. & Jacobs, J.A. (1994). "A Cross-National Analysis of the Wages of Part-Time

Workers: Evidence from the United States, United Kingdom, Canada and Australia." LIS Working Paper # 113. Syracuse, NY: University of Syracuse

Gornick, J.C. (1998). "Income Maintenance and Employment Supports for Former Welfare Recipients: The United States in Cross-National Perspective." Prepared for the National Governor's Association, November 8, 1998

Gornick, J.C., Meyers, M.K. & Ross, K.E. (1997). "Supporting the Employment of Mothers: Policy Variation Across Fourteen Welfare States." Journal of European Social Policy, 7(1), pp.45-70

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

Hammermesh, D.S. (1982). Jobless Pay and the Economy. Baltimore, MD: Johns Hopkins University Press

Hansen, G.D. & Imrohoroglu, A. (1992). "The Role of Unemployment Insurance in an Economy with Liquidity Constraints and Moral Hazard." Journal of Political Economy, 100(1), pp.118-142

Hatcher, C. (1999). "A Life-Course Approach to Measuring Household Wealth." Journal of Consumer Affairs, 33(2), pp.358-379

Heintze, T. (1999). "Unemployment Compensations Programs' Effect on the Employment of

Young Men: A Cross-National Comparison of Canadian, British and American Unemployment Policies." Presentation at 1999 Association for Public Policy Analysis and Management Meeting, Washington, DC. November 4-6, 1999

Human Resource Department of Canada (1996). The New Employment Insurance System. Ottawa, CN: Author

Inland Revenue (1999). National Insurance for Employees. London, UK: Author

Killingsworth, M.R. (1983). Labor Supply. Cambridge, UK: Cambridge University Press

Kloosterman, R.C. (1994). "Three Worlds of Welfare Capitalism? The Welfare State and the Post-Industrial Trajectory in the Netherlands after 1980." West European Politics, 17(4), pp.166-189.

Korpi, W. & Palme, J. (1998). "The Paradox of Redistribution and Strategies of Equality: Welfare State Institutions, Inequality and Poverty in the Western Countries." LIS Working Paper #174. Syracuse, NY: Syracuse University

Kost, K.A. (1996). "A Man Without a Job is a Dead Man: The Meaning of Work and Welfare in the Lives of Young Men." Institute for Research on Poverty Discussion Papers: DP#1112-96. Madison, WI: University of Wisconsin-Madison

Lancaster, T & Nickell, S.J. (1980). "The Analysis of Re-Employment Probabilities for the Unemployed." Journal of the Royal Statistical Society, 143, pp.141-152

Lipset, M.S. (1989). "American Exceptionalism Reaffirmed." Unpublished

Marquardt, D.W. (1963). "An Algorithm for Least Squares Estimation of Nonlinear Parameters." Journal of the Society for Industrial and Applied Mathematics, 11, pp.431-441

Meyer, B.D. (1990). "Unemployment Insurance and Unemployment Spells." Econometrica, 58(4), pp.757-782

Meyers, M.K., Heintze, T. & Wolf, D.A. (2002). "Child Care Subsidies and the Employment of Welfare Recipients." Demography, 39(1), pp. 165-180

Milbourne, R.D., Purvis, D.D. & Scoones, W.D. (1991). "Unemployment Insurance and Unemployment Dynamics." Canadian Journal of Economics, 24(4), pp.804-826

Moffitt, R. & Nicholson, W. (1982). "The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits." The Review of Economics and Statistics, 64, pp.1-11

Moffitt, R. (1983). "An Economic Model of Welfare Stigma." The American Economic Review. 73(4), pp.1023-1035

Moffitt, R. (1992). "Incentive Effects of the U.S. Welfare System: A Review." Journal of Economic Literature, 30(March), pp.1-61

Mortensen, D.T. (1977). "Unemployment Insurance and Job Search Decisions." Industrial and Labor Relations Review, 30, pp.505-517

Murphy, K.M. & Topel, R.H. (1985). "Estimation and Inference in Two-Step Econometric Models." Journal of Business and Economic Statistics, 3(4), pp.370-379

National Council of Welfare (1998). Welfare Incomes 1996: A Report by the National Council of Welfare. Ottawa, CN: Author

Naylor, J.C. & Smith, A.F.M. (1982). "Applications of a Method for the Efficient Computation of Posterior Distributions." Applied Statistics, 31(3), pp.214-225

Organization for Economic Cooperation and Development (1982). The OECD List of Social Indicators. Paris, FR: Author

Organization for Economic Cooperation and Development (1994). "Unemployment and Related Welfare Benefits." In The OECD Jobs Study: Evidence and Explanations; Part II: The Adjustment Potential of the Labor Market (pp.171-237). Paris, FR: Author

Organization for Economic Cooperation and Development (1996). Thematic Review of Unemployment Benefits. Paris, FR: Author

Organization for Economic Cooperation and Development (1998). OECD Composite Leading Indicators. Paris FR: Author

Pissarides, C.A. (1991). "Real Wages and Unemployment in Australia." Economica, 58(February), pp.35-55

Portes, A. & Zhou, M. (1996). "Self-Employment and the Earnings of Immigrants." American Sociological Review, 61, pp.219-230

Pressman, S. (1995). "The Gender Poverty Gap in Developed Countries: Causes and Cures." LIS Working Paper #135. Syracuse, NY: Syracuse University

Rubin, D.B. (1987). Multiple Imputation for Nonresponse in Surveys. New York, NY: John Wiley & Sons

Savoie, D.J. (1994). Thatcher, Reagan and Mulroney: In Search of a New Bureaucracy. Pittsburgh, PA: University of Pittsburgh Press

`

Schafer, J.L. (1997). Analysis of Incomplete Multivariate Data. London, UK: Chapman and Hall

Smeeding, T. & Gottschalk, P. (1996). "The International Evidence on Income Distribution in Modern Economies: Where do We Stand?" LIS Working Paper #137, Syracuse NY: Syracuse University

Smeeding, T. (1995). "Evaluating Inequality: An Overview" Center for Demography and Economics of Aging Seminar Series Workshop Paper #11. Syracuse, NY: Syracuse University

Stata Press (1997). Stata Reference Manual, Release 5. College Station, TX: Author

Steck, G.P. (1958). "A Table for Computing Trivariate Normal Probabilities." Annuals of Mathematical Statistics, 29(3), pp.780-800

Steen, N.M., Byrne, G.D. & Gelbard, E.M. (1969). "Gaussian Quadratures for the Integrals: $_0\int^{\infty} \exp\{-x^2\} f(x)\,dx$ and $_0\int^{b} \exp\{-x^2\} f(x)\,dx$" Mathematics of Computation, 23(107), pp.661-671

Stroud, A.H. & Secrest, D. (1966). Gaussian Quadrature Formulas. Englewood Cliffs, NJ: Prentice-Hall Inc.

Sullivan, D.H. & Smeeding, T. (1997). "Educational Attainment and Earnings Inequality in Eight Nations." LIS Working Paper # 164. Syracuse, NY: Syracuse University

Tannery, F. (1983). "Search Effort and Unemployment Insurance Reconsidered." Journal of Human Resources, 18, pp.432-440

Trattner, W.I. (1994). From Poor Law to Welfare State: A History of Social Welfare in America (Fifth Edition). New York, NY: The Free Press

U.S. Department of Labor (1999). Comparison of States' Unemployment Insurance Laws. Washington, DC: Author

U.S. House of Representatives (1996). 1996 Green Book. Washington, DC: Author

Wadsworth, J. (1991). "Unemployment Benefits and Search Effort in the UK Labour Market." Economica, 58(February), pp.17-34

Wandner, S.A. & Stengle, T. (1997). "Unemployment Insurance: Measuring Who Receives It." Monthly Labor Review, 120(July), pp.15-24

Wolf, D.A. (2000). "Microsimulation, Population Projection and Interval Estimation". Presentation at Population Research Institute, Pennsylvania State University, April 11, State College, PA

Yang, Z. (1998). "Comparison of Two Gaussian Quadratures in Multivariate Normal Integrals." Chinese Journal of Numerical Mathematics and Applications, 20(1), pp.59-69

Table I-1: Comparison of Indicators for Esping-Anderson Welfare States

	Canada	United Kingdom	United States	Sweden	Germany
% Labor Force in Services Sector [a] (1990)	72.2 %	69.7 %	72.0 %	67.9 %	57.6 %
% Labor Force Female [a] (1990)	44.7 %	43.6 %	45.4 %	47.8 %	40.7 %
Child Orientated Policies Enabling Female Labor Participation [b]: (1994/5)					
Under School Age	32.4 (9)	21.6 (12)	17.1 (14)	61.9 (3)	34.1 (8)
Under Age Three	34.7 (10)	22.0 (12)	13.6 (14)	62.3 (3)	36.2 (8)
Age Three to Five	30.1 (9)	21.3 (11)	20.7 (13)	61.6 (4)	31.9 (8)

[a] Source: Godbout (1993)
[b] Source: Gornick, Meyers & Ross (1997)

Table I-2: Economic and Poverty Statistics for Canada, the United Kingdom and the United States (1994-1995)

	Canada	United Kingdom	United States
% Males (18-65) in Labor Force	84.8 %	85.5 %	87.7 %
% Male Labor Force (18-65), Unemployed	9.7 %	8.8 %	5.3 %
Poverty Rate	10.62	12.44	18.76
P90/P10 Inequality Measure	3.88	4.54	6.52

Source: LIS Wave IV, Author's Calculations, N=9,955 weighted to population

Table I-3: Means and Standard Deviations of Variables Used in Analysis

	Canada	United Kingdom	United States
Demographic Variables			
Age	21.62	21.55	21.76
	(2.31)	(2.33)	(2.32)
% Married or Cohabiting	16.91 %	24.10 %	21.51 %
% of Men with Own Household	28.19 %	30.45 %	37.18 %
% Self Employed	5.08 %	5.87 %	3.16 %
Education Level 1	24.57 %	49.07 %	21.63 %
Education Level 2	27.74 %	25.93 %	35.72 %
Education Level 3	41.86 %	18.28 %	32.82 %
Education Level 4	5.82 %	6.72 %	9.84 %
Housing Tenure (Years)	1.92	4.38	1.42
	(0.84)	(1.81)	(0.49)
Number of Children in Household	0.63	0.69	0.92
	(0.97)	(1.07)	(1.15)
Age of Youngest Child in Household	3.82	3.69	3.83
	(6.12)	(5.97)	(5.74)
Number of Earners in Household	2.82	2.15	2.63
	(1.19)	(1.13)	(1.15)
Other Equivalent Income	$32,165.04	$27,219.84	$30,307.10
	($24,499.97)	($25,930.90)	($28,339.26)
UC Policy Variables			
Maximum Weekly Insurance Benefit	$344.17	$65.94	$290.48
	($0.00)	($16.83)	($58.70)
Maximum Weekly Assistance Benefit	$102.27	$65.94	$68.37
	($27.16)	($16.83)	($10.15)
Maximum Insurance Duration (Weeks)	37.76	52	26.18
	(5.43)	(0.00)	(0.82)
Assistance Benefit Disregard	$21.27	$14.93	$0.00
	($11.02)	($0.00)	($0.00)
Insurance Benefit Disregard	$79.24	$7.42	$70.67
	($18.70)	($0.00)	($47.69)
National/Regional Variables			
% Lives in Low Population Density Area	25.62 %	25.91 %	20.5 %
Regional Unemployment Rate	10.41 %	8.66 %	6.16 %
	(2.11)	(1.54)	(1.32)
Male Median Wage	$14.43	$10.97	$13.13
	($1.17)	($1.35)	($1.52)
Tax Rate	0.15	0.193	0.15
	(0.08)	(0.10)	(0.07)

Note: Standard deviations in parentheses, N=9,955 weighted to population

Table I-4: Comparison of Labor Statistics and Unemployment Compensation for Young Men (18-25) and Older Men (26-65), 1994-1995

	Canada		United Kingdom		United States	
	Young Men	**Older Men**	**Young Men**	**Older Men**	**Young Men**	**Older Men**
Labor Force:						
% In Labor Force[a]	71.49%	86.6%	83.34%	85.81%	78.58%	89.15%
% Labor Force Unemployed[a]	17.08%	8.73%	15.55%	7.89%	11.29%	4.44%
UC Receipt:						
% Labor Force Receiving Unemployment Insurance[b]	15.52%	13.05%	2.78%	0.99%	5.52%	6.37%
% Labor Force Receiving Unemployment Assistance[b]	10.41%	7.24%	15.68%	7.9%	6.61%	11.88%
% Labor Force Receiving Unemployment Compensation[b]	23.58%	18.65%	17.61%	8.66%	11.32%	16.93%
% of Unemployed Receiving Unemployment Compensation[c]	45.43%	66.91%	86.11%	76.81%	33.27%	61.91%
UC Benefits:						
Average Benefit in Past Year	$4,755.19	$5,793.52	$4,682.22	$6,006.86	$1,670.20	$2,008.97
Average Weekly Benefit	$251.28	$275.54	$168.07	$213.58	$177.90	$236.57
Average Weeks Unemployed[d]	24.31	25.61	31.63	29.91	18.57	19.6
Labor Earnings and Income:						
Median Wage[e]	$9.48	$16.03	$6.64	$11.66	$7.38	$14.72
Average Hours Worked Per Week[e]	33.98	42.93	38.85	43.93	37.22	44.82
Men's Median Household Income	$35,821.67	$38,377.5	$26,914.27	$28,298.79	$30,528.57	$40,433.67
Average UC/Household Income	5.31	4.49	5.92	3.59	2.24	1.67

[a] Measured at present time
[b] Measured in past year
[c] Measure of those unemployed at current time who received unemployment compensation in past year
[d] Measured in the past year among those who experience unemployment
[e] Current measure for those working at current time

Source: LIS Wave IV, Author's Calculations, N=9,955 weighted to population

Table I-5: Probit Used to Predict Insurance Eligibility for Young Men

Variable	Coefficient	Standard Error[1]	
Programmatic Variables			
Maximum Insurance Benefit	0.00178	0.00100	
Maximum Duration of UI Receipt	0.02480	0.01738	
Minimum Wage to Receive UI	-0.00001	0.00007	
Minimum Work Time to Receive UI	0.01225	0.01166	
Demographic Variables			
Education Level 1	-0.17102	0.24102	
Education Level 2	-0.08062	0.24166	
Education Level 3	-0.11480	0.24017	
Self Employed	-0.62338	0.25375	*
National/Regional Variables			
Canadian	0.46695	0.20254	*
British	-1.22611	0.61869	*
Regional Unemployment Rate	0.04327	0.04074	
Employment Categories			
Management	-0.05678	0.41183	
Professional	0.16285	0.29539	
Technical	0.30191	0.32881	
Clerk	-0.1523	0.26526	
Service	0.65166	0.29792	*
Agriculture	0.22606	0.30852	
Craft	0.42000	0.27652	
Operators	0.32431	0.27559	
Unskilled	0.66770	0.35570	
Armed Forces	0.39023	0.40506	
Inverse mills ratio (Known Eligibility)	-1.46854	0.18977	***
Constant	-1.58011	0.68527	*
Log Likelihood	-851.39		
Chi^2	273.18		
N	1979		
Pseudo R^2	0.2466		

* $p < 0.05$ ** $p<0.01$ *** $p<0.001$ (two tailed)

[1]Standard errors are uncorrected for presence of variable from previous stage (inverse mills ratio), implying that they are understated.

Table I-6: Probit Used to Predict Whether Young Men are in School

Variable	Coefficient	Standard Error	
Demographic Variables			
Married or Cohabiting	-0.90080	0.40254	*
Age	-0.18210	0.03617	***
Education Level 1	-0.17970	0.36502	
Education Level 2	-0.41112	0.35487	
Education Level 3	-0.21700	0.34687	
Young Man Heads Own Household	-0.53097	0.22653	*
Number of Children in Household	-0.03123	0.07701	
Age of Youngest Child in Household	0.02379	0.01319	
Housing Tenure	-0.17590	0.09495	
Non-Worker	1.79360	0.29396	***
Number of Earners in Household	-0.47215	0.08985	***
Other Equivalent Income in Household	0.00001	0.00019	
Regional Variables			
Lives in Rural Area	-0.35897	0.16234	*
Regional Unemployment Rate	-0.01765	0.03372	
Constant	3.77131	1.05731	***
Log Likelihood	-315.15		
Chi^2	190.25		
N	3626		
Pseudo R^2	0.3212		

* $p < 0.05$ ** $p<0.01$ *** $p<0.001$ (two tailed)

Table I-7: Means and Standard Deviations of Variables After Multiple Imputation Completed

Variable	N of "Knowns"	Among "Knowns"	Full Sample (includes imputed values), N = 9955				
			Run 1	Run 2	Run 3	Run 4	Run 5
% Eligible for Insurance	1979	0.475	0.531	0.565	0.560	0.551	0.562
		(0.500)	(0.499)	(0.496)	(0.496)	(0.497)	(0.496)
% In School	3626	0.023	0.039	0.042	0.045	0.040	0.034
		(0.151)	(0.192)	(0.201)	(0.207)	(0.197)	(0.181)

Note: Standard deviations in parentheses

Table I-8: Regression Equation Used to Predict Hourly Wage

Variable	Coefficient	Standard Error[1]	
Demographic Variables			
Education Level 1	-4.50175	1.15633	***
Education Level 2	-4.48411	0.98928	***
Education Level 3	-1.56841	0.96499	
Age	0.45490	0.13903	***
Self Employed	-0.94214	0.95689	
Canadian	-0.19368	1.03700	
British	-4.94994	0.82984	***
National/Regional Variables			
Lives in a Rural Area	-2.19726	0.64043	***
Regional Unemployment Rate	-0.66978	0.19890	***
Male Median Wage	0.12258	0.19595	
Inverse Mills Ratio (Workers)	-15.31542	0.17843	***
Constant	36.71957	5.72049	***
N	8591		
R^2	0.0786		

* $p < 0.05$ ** $p<0.01$ *** $p<0.001$ (two tailed)

[1]Standard errors are corrected for presence of variable from previous stage (inverse mills ratio).

Table I-9: Trivariate Probit of Work, Assistance Receipt and Insurance Receipt (Model 1)

Variables	Work			Assistance Receipt			Insurance Receipt		
	Coefficient	Error		Coefficient	Error		Coefficient	Error	
Programmatic Variables									
Maximum Weekly Assistance Benefit	-0.00164	0.00079	*	0.00472	0.00086	***	0.00467	0.00113	***
Maximum Weekly Insurance Benefit	-0.00046	0.00037		-0.00181	0.00068	**	-0.00016	0.00073	
Maximum Duration of UI Receipt (Weeks)	0.00841	0.00532		-0.01628	0.01011	a	-0.00106	0.00787	
Demographic Variables									
In School	-0.40591	0.08200	***	-0.16993	0.09993	a	-0.06587	0.17826	
Predicted Wage	-0.03534	0.05447		-0.00377	0.24197		0.03787	0.09483	
Education Level 1	-0.65697	0.32329	*	0.06266	1.52521		0.80184	0.57760	
Education Level 2	-0.3738	0.29035		-0.01225	1.37545		0.65213	0.53663	
Education Level 3	-0.11367	0.11346		-0.05450	0.47669		0.29281	0.20781	
Married or Co-Habiting	0.14447	0.04921	**	-0.07958	0.05828		-0.06387	0.06622	
Age	0.02524	0.03214		0.00595	0.14167		0.08518	0.05477	
Young Man Heads Own Household	0.13283	0.04473	**	0.07302	0.05559		0.12011	0.06933	a
Number of Children in Household	0.03322	0.02253		0.10651	0.03630	**	0.08722	0.04427	*
Age of Youngest Child in Household	-0.00503	0.00337		0.00024	0.00416		-0.00124	0.00525	
Housing Tenure	0.02473	0.02212		0.11080	0.02582	***	0.04295	0.03310	
Self Employed	0.31753	0.09048	***	0.66988	0.27452	*	-0.06169	0.19599	
Number of Earners in Household	0.13486	0.01775	***	-0.07013	0.02346	**	-0.10909	0.02738	***
Other Equivalent Income (in $1,000s)	-0.15000	0.05000	**	-0.06000	0.06000		-0.08000	0.08000	
Regional/National Variables									
Canadian	-0.06534	0.06198		-0.63736	0.10553	***	-0.05963	0.09655	
British	-0.41978	0.29502		0.03635	1.28113		-0.21853	0.59933	
Lives in Rural Area	-0.20560	0.13209		-0.02951	0.58358		0.36142	0.23068	
Regional Unemployment	-0.08714	0.04886	a	-0.01277	0.22618		0.09270	0.08891	
Tax Rate	2.21161	0.25239	***	-1.91216	0.50787		-1.36967	0.41106	***
Constant	2.45489	2.35423		0.92110	10.74441		-5.52264	4.07580	
ρ (Work, Assistance)	-0.33096	0.02734	***						
ρ (Work, Insurance)	-0.32282	0.02952	***						
ρ (Assistance, Insurance)	0.37291	0.05452	***						
Log Likelihood	-11266.729								
N	9955								
Chi^2	2799.24								
Pseudo R^2	0.1105								

+ p < 0.1 * p < 0.05 ** p<0.01 *** p<0.001

Table I-10: Trivariate Probit of Work, Assistance Receipt and Insurance Receipt (Model 2)

Variables	Work			Assistance Receipt			Insurance Receipt		
	Coefficient	Error		Coefficient	Error		Coefficient	Error	
Programmatic Variables									
Maximum Weekly Assistance Benefit	-0.00150	0.00078	a	0.00485	0.00089	***	0.00255	0.00113	*
Maximum Weekly Insurance Benefit	-0.00094	0.00040	*	-0.00175	0.00112		0.00303	0.00085	***
Maximum Duration of UI Receipt (Weeks)	0.00359	0.00673		-0.01982	0.01263		0.02939	0.00891	**
Weekly Assistance Benefit Disregard	0.00190	0.00279		-0.00135	0.00401		-0.01194	0.00339	***
Weekly Insurance Benefit Disregard	0.00148	0.00047	**	0.00039	0.00049		-0.01171	0.00070	***
Demographic Variables									
In School	-0.41583	0.07652	***	-0.19322	0.09963	a	-0.04049	0.14102	
Predicted Wage	-0.08689	0.06514		0.01016	0.25491		0.51885	0.25482	*
Education Level 1	-0.95488	0.39863	*	0.16455	1.61245		3.46519	1.76559	*
Education Level 2	-0.63889	0.34942	a	0.09630	1.44715		3.06927	1.47041	*
Education Level 3	-0.19690	0.12871		0.01265	0.49206		1.08800	0.44943	*
Married or Co-Habiting	0.14864	0.04895	**	-0.06383	0.06565		-0.02589	0.06509	
Age	0.05445	0.03896		-0.00211	0.14910		-0.18484	0.16203	
Young Man Heads Own Household	0.14355	0.04454	**	0.07251	0.05768		0.09467	0.06913	
Number of Children in Household	0.02840	0.02242		0.10706	0.03548	**	0.10209	0.04246	*
Age of Youngest Child in Household	-0.00513	0.00336		0.00127	0.00399		-0.00070	0.00536	
Housing Tenure	0.02745	0.02221		0.11319	0.02727	***	0.03303	0.03186	
Self Employed	0.33511	0.11042	**	0.60515	0.30789	*	-0.21877	0.51041	
Number of Earners in Household	0.13895	0.01765	***	-0.07570	0.02175	***	-0.09642	0.02602	***
Other Equivalent Income (in $1,000s)	-0.15000	0.05000	**	-0.04000	0.07000		-0.10000	0.07000	
Regional/National Variables									
Canadian	-0.09919	0.07079		-0.61853	0.10828	***	0.06296	0.09906	
British	-0.62404	0.32532	a	0.22639	1.33106		1.57262	1.20216	
Lives in Rural Area	-0.33049	0.15775	*	-0.00548	0.61399		1.45610	0.62407	*
Regional Unemployment Rate	-0.12143	0.05968	***	0.00473	0.23679		0.43905	0.25950	a
Tax Rate	2.28694	0.25280	***	-2.09280	0.44703	***	-0.89705	0.39033	*
Constant	4.75757	2.88069	a	0.29311	11.37235		-26.5975	12.16755	*
ρ (Work, Assistance)	-0.31241	0.02560	***						
ρ (Work, Insurance)	-0.28489	0.02954	***						
ρ (Assistance, Insurance)	0.35240	0.05147	***						
Log Likelihood	-11020.818								
N	9955								
Chi^2	3255.404								
Pseudo R^2	0.1287								

+ $p < 0.1$ * $p < 0.05$ ** $p<0.01$ *** $p<0.001$

Table I-11: Predicted Probability of Work at Selected Values of Insurance Benefit Duration and Insurance Benefit Disregard

Model 1		Model 2	
Maximum Duration of Insurance Benefits (Weeks)	**Probability of Work for Individual at Mean**	**Insurance Benefit Disregard (U.S. $)**	**Probability of Work for Individual at Mean**
20	0.851	40	0.870
25	0.861	50	0.873
31 (Mean)	0.871	58 (Mean)	0.876
35	0.879	70	0.879
40	0.887	80	0.882

Sample probability of working: 87.7 %