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SOCIAL ASSISTANCE AND LABOUR SUPPLY

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Abstract

The longitudinal, 1988-89, LMAS makes it possible to study the interaction between social assistance and labour supply while allowing for substantial time dependence. Tobit equations for hours worked on and off social assistance, which allow for endogenous selection of social assistance status, are estimated using FIML. Small, but statistically significant, effects for a social assistance benefit variable are obtained, particularly for females. A claw-back variable is not consistently useful. The wage rate has a consistently negative coefficient in the selection equations and is important in the Tobit equations. The paper attempts to reconcile results in earlier literature.

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1. Introduction

The social assistance program in Canada[1] has grown in size and, in light of the difficult financial circumstances faced by all levels of government, it has attracted the attention of policy makers. Some provinces have implemented, or are planning, reforms to their systems. Where reforms have already proceeded, this appears to have occurred because of political imperatives rather than in response to findings from studies of the determinants of participation in social assistance programs and in the labour market. It is not clear whether the time series reductions in the Ontario social assistance caseload[2] are due to recent reductions in the generosity of the program, administrative changes, the quickening pace of economic activity, or all of these forces. Questions such as these must be posed in the context of a model that is capable of addressing simultaneously the social assistance outcome and labour market involvement for, with sufficient hours of work and/or a high-enough wage rate, an individual would not be eligible for assistance. Using cross-sectional evidence, this paper takes a step in this direction by considering the role of the

[1] Under the Canada Assistance Plan (CAP) Act the federal government assumes about half the cost of social assistance. Provinces do have substantial powers in the design, administration and delivery of social assistance. While most provinces have a unified system, others such as Nova Scotia, Manitoba and Ontario operate systems where municipalities as well as the province have certain responsibilities - depending on the characteristics of the claimant. Eligibility is based on a needs test and takes household income into account. Reference to Canada's social assistance program, therefore, should be understood as reference to an exceedingly complex set of eligibility rules, benefit structures, earnings exemptions, treatments of fixed and liquid assets, children and disability. In 1996, CAP was replaced by the block fund in the Canada Health and Social Transfer.

[2] For details, see note 36 below.

parameters of the social assistance program in shaping program participation and labour market involvement decisions. Section 2 reviews Canadian work on this topic. While a number of studies have been undertaken, these tend to focus on the behaviour of a particular group (mainly lone mothers and immigrants), aspect of activity (the social assistance decision), or experience in a particular province.

The paper treats social assistance participation and labour market involvement simultaneously and extends the analysis to the widest sample that can be considered on a Canada-wide basis. The behaviour of groups, such as lone mothers and immigrants which are of particular public policy interest, is considered. Unlike earlier work, which has tended to study only women, this paper considers observations pertaining to both genders and a variety of family units - e.g. single individuals, lone parents and individuals who have a spouse present and may, or may not, have children. The cross-sectional nature of the data used makes it possible to condition the analysis on parameters of social assistance programs which differ by province, thereby injecting considerable cross-sectional variability into the statistical work; further variance in the social assistance variables is induced as the individual's disability status and number of children are taken into account. An advantage of the data base used over other cross-sectional data is that it is possible to condition on the individual's participation in social assistance and in the labour force during the year prior to the one under study. An attempt is also made to consider the extent to which divergent results in the existing, exclusively cross-sectional, literature can be reconciled. Section 3 outlines the conceptual and econometric framework adopted while section 4 considers data issues. Section 5 contains the results obtained and section 6 presents concluding observations.

2. Earlier Literature

Interest by economists in Canada's program of social assistance is relatively recent. Allen (1993) explores the effects of the basic social assistance and liquid asset exemption parameters in provincial programs on welfare participation, family structure and labour force participation among low-income women. The model treats these decisions as independent and Allen (1993) concludes that the elasticity of labour force participation with respect to social assistance benefits is small while the estimated social assistance participation and family structure elasticities are more substantial. Using the Labour Market Activity Survey (LMAS), Charette and Meng (1994) find that while social assistance benefit levels and especially earned income exemptions may play a significant role in the participation decision of female heads of households, the effect of the program claw-back rate on earnings above the exemption level is less clear. Dooley (1994) concludes that the level of social assistance is a significant determinant of program participation for lone mothers. Baker and Benjamin (1995a) consider the receipt of social assistance among census families using the 1986 and 1991 Surveys of Consumer Finances (SCF). They are particularly interested in the behaviour of immigrant families, defined here as those whose head was born outside Canada. They find that, on entry, immigrant cohorts are less likely to be on social assistance than the Canadian-born, that, again on entry, more recent cohorts are more likely to be on social assistance than earlier cohorts of immigrants and that the probability of social assistance receipt increases over time spent in Canada for all

cohorts.[3] The 1994 CERF/BCMSS Workshop on Labour Markets and Income Support examined aspects of welfare behaviour often using data taken from the British Columbia Income Assistance Program. Particular attention is paid to the duration and duration dependence of spells on social assistance. Bailey (1994) evaluates the impact of the B.C. Enhanced Earnings Exemption on labour supply and on exit rates from the Income Assistance Program by family type. He concludes that, while more generous earnings exemptions may have increased the earnings and labour supply of social assistance recipients, the general effect of the new exemption levels on exit from the program is unclear. Evidence of a small retention effect was found for lone parent families but not for other family structures. The dynamics of social assistance dependence in British Columbia is also examined in Barrett and Cragg (1998) and Barrett (1996). The latter finds that exit from social assistance is negatively related to the generosity of the social assistance program and a similar result appears to hold for Quebec in Fortin, Lacroix and Thibault (1997). Stewart and Dooley (1997) study the duration of welfare spells using administrative data from Ontario. Card and Robins (1996) present a progress report on the Self-Sufficiency Project. Thus, the existing literature tends to suggest that the generosity of the social assistance system has some impact on both the probability of program participation and the duration of spells on social assistance but it is generally silent on the issue of whether program parameters have an effect on the supply of labour.

Christofides, Stengos and Swidinsky (1997a) (CSS) appears to be the only Canadian study of the interaction between participation in the labour force and

[3] Further discussion on the use of social assistance by immigrants can be found in Baker and Benjamin (1995b) and Lui-Gurr (1995).

social assistance. It adopts a Bivariate Probit model of participation in the labour force and in social assistance and uses it to examine selection-adjusted hours equations for individuals on and off social assistance. This study, which considers separately single males and females and lone fathers and mothers, concludes that the parameters of the social assistance program do not appear to exert any influence on the decisions of these four groups. Instead, personal characteristics (such as disability and the number of children present) and labour market conditions (such as the wage rate and regional indicators) are seen to be important. These results raise the question of whether they stem from the integrated framework adopted, their particular treatment of sample selection, or the size of the samples used and they call for further investigation of the interaction between use of social assistance and labour supply.

CSS and this paper both focus on this interaction but they differ in important model, sample and variable-used respects. By specifying a bivariate probit model for the social assistance and labour force participation decisions, CSS implicitly split the behaviour analyzed by the Tobit labour supply equation in the present paper into the labour force participation decision on the one hand and the desire to supply hours of work conditional on participation on the other. In principle, this separation permits the analyses of participation and extent of work to proceed along different lines and does not impose the constraint, implicit in the Tobit approach, that variables must affect the two aspects of behaviour in proportion and with the same sign.[4] In practice, this advantage is difficult to exploit given

[4]The effect of a continuous variable on the *unconditional* expected value of hours supplied is proportional to its effect on the probability of labour force participation, the proportionality factor being a function of the ratio of the standard normal density and cumulative density

that available data do not specifically pertain to the participation (as distinct from the supply of hours) decision. Moreover, it requires the estimation of hours to proceed using Heckman-type corrections for sample selection. It is well-known that these corrections may lead to results which are sample dependent[5] and it is preferable to bypass the sample selection problem altogether. The model used in this paper does this and offers a sensitivity analysis by providing estimates of the effects of program parameters based on a different framework.

The model in CSS culminates in equations for the hours of work by individuals who do work and are either on or off social assistance. These equations must be estimated using small sub-samples of individuals on social assistance which range from 24 to 243 observations for lone fathers and mothers respectively. In the present paper, the labour market involvement of individuals on and off social assistance is examined using endogenously selected Tobit equations which are based on considerably larger samples. While men and women are considered separately, possible differences in the behaviour of single individuals and lone parents are modelled through a marital status indicator and variables signalling the presence of children. A substantial number of additional observations are possible if the sample is expanded to include, as some earlier studies have, individuals who have a spouse present. The addition of this group raises issues concerning conditioning variables which are addressed below. These objectives are pursued using the longitudinal 1988-89 LMAS which, working with the 1989 data, makes it possible to take into account an individual's social assistance and work status during 1988.

functions. The ratio is necessarily positive and hence the two effects must have the same sign.

[5] Baker *et al* (1995) also note this point. Christofides and Liu (1999) review and illustrate the advantages of recently proposed alternatives to the Heckman approach which are less sample-specific and do not rely on normality.

These variables sharpen the empirical analysis dramatically yet still 'leave room' for other important forces. In CSS, only past social assistance use was taken into account and then only in the hours equations for non-welfare participants, where it was generally not significant. Thus, this paper offers a very different examination of the interaction between participation in the social assistance program and labour force involvement than CSS.[6]

3. Social Assistance and Labour Supply

A number of US studies in this area[7] adopt a structural approach and assume that individuals maximize a utility function which depends on income and leisure subject to a budget constraint. This constraint incorporates labour and non-labour income as well as possible income from social assistance.[8] Utility is maximized by choice of hours of work and, consequently, labour income and the individual accepts social assistance if, *ceteris paribus*, the value of the indirect utility function for a participant exceeds that for a non-participant. Possible stigma effects of participation in social assistance are taken into account. Moffitt's (1983) econometric implementation of this approach assumes a linear hours of work equation which, through the integrability conditions, implies a particular utility function. The thrust of the approach is to provide estimates of the utility function parameters.

While accepting this broad approach, our empirical implementation is not

[6] Detailed differences between the two papers are noted below - e.g. the last paragraph in section 3 and note 30.

[7] See Basi (1990), Blank (1985) and Moffitt (1983) for examples of this approach and further references. Recent UK work on related issues includes Bingley *et al* (1995) and Bingley and Walker (1997).

[8] It is assumed that there are no eligibility constraints based on other factors such as wealth.

focused on the estimation of preference parameters but, rather, the program participation and labour supply outcomes themselves. This more structurally agnostic specification[9] specifies a Probit social assistance participation equation. The labour force participation and hours behaviour of individuals on and off social assistance are specified as Tobit equations. More specifically, the program participation equation is

$$P_j^* = X_j\alpha + u_j \tag{1}$$

where the actual participation indicator for the jth individual is given by

$$P_j = 1 \text{ if } P_j^* > 0 \text{ and} \tag{2}$$

$$P_j = 0 \text{ if } P_j^* \leq 0 \tag{3}$$

and the X_j are $1 \times p$ individual-specific vectors of variables, α is a $p \times 1$ vector of constant coefficients and u_j is a standard normal variable. The hours-of-work equations are given by

$$H_j^{i*} = Z_j^i\beta^i + e_j^i \tag{4}$$

where the superscript i indexes the social assistance states 0 = non-participation and 1 = participation, Z_j is a $1 \times h$ vector of individual-specific variables, β is

[9]Social activists argue that participation in social assistance is not a matter of choice but a measure of last resort arising from personal ciscumstances such as disability or labour market rationing. The state of the labour market cannot readily be incorporated in the standard optimising framework. In the empirical work below, we condition on regional dummies which, among other forces, capture the very diverse labour market states in the regions of Canada.

an $h \times 1$ vector of constant coefficients and the random variables $e^{\mathbf{0}}$ and e^{1} are normally distributed with zero means, constant variances $\sigma^{\mathbf{00}}$ and σ^{11} and covariances $\rho^{\mathbf{0}u}$ and ρ^{1u}; the covariance $\rho^{\mathbf{01}}$ is assumed equal to zero because the two social assistance states are mutually exclusive. The observed number of hours is given by

$$H_j^i = H_j^{i*} \text{ if } H_j^{i*} > 0 \text{ and} \tag{5}$$

$$H_j^i = 0 \text{ if } H_j^{i*} \leq 0. \tag{6}$$

The Probit participation equation and the Tobit hours-of-work equations are estimated simultaneously using full information maximum likelihood (FIML) procedures, though we also report estimates based on separate Probit and Tobit equations. The model as specified is much closer to the econometric version of Moffitt (1983) than CSS. An advantage of the FIML approach is that the effect of any explanatory variable on, say, the social assistance decision is estimated taking into account its possible effect on the labour supply decision. In addition, when the equations of the model are specified correctly and estimated simultaneously, the sample-selection problem involved in analyzing individuals on, or off, social assistance does not arise. Moreover, the sample selection problem involved in analyzing individuals with positive hours of work only does not arise either because the Tobit likelihood function consists of the sum of the Probit (choice of being in the labour-force or not) and Truncated Regression (analysis of the positive hours sample) likelihoods and hence takes into account this second selection problem. Standard selection-correction procedures may produce results which, as

noted above, may be sample-dependent and it is desirable, as our present approach does, to avoid the selection problem from the beginning.

4. Sample Construction, Variables and Sources

Data for this paper are for 1989 and are drawn from the 1988-89 Longitudinal LMAS. Beginning with the full sample, observations with any of the following characteristics were omitted: (i) Full-time students,[10] (ii) individuals with an unstated or unknown province of residence[11], and (iii) persons over 64 years of age in 1988. Two main samples are considered. The first sample, which consists of 14,101 observations, excludes individuals who have a spouse present.[12] The reason for focusing on this particular sample is that the LMAS reports information on individuals only while the decisions to apply and provide social assistance are based on family characteristics. Thus, an individual who has attributes that might normally be associated with social assistance receipt may, in fact, not be a program participant because of the income earned by a spouse. This issue is irrelevant in samples where a spouse is not present.[13] A possible problem with rigorous adherence to this principle of sample construction may be the size of the implied

[10] Variable Q152ST89 identifies individuals who were full-time students during 1989.

[11] This information is needed for the calculation of the parameters of the social assistance program and in order to control for regional labour market tightness.

[12] In the LMAS variable HHCHAR89, category 1 consists of individuals with no spouse and no children, while category 4 comprises individuals with no spouse who have children present. Individuals in these two categories make up the main sample used in this paper. Individuals who have a spouse present are analysed in the second sample below and they include those without children (category 3) and those with children (category 4) present.

[13] On the exogeneity of family structure see Dooley, Lefebvre and Merrigan (1997).

sample. Table 1 shows that the first sample consists of 6,932 males and 7,169 females[14] of whom 611 and 1135, respectively, received social assistance.[15] The number of individuals who received social assistance while in the labour force[16] was 305 males and 558 females. A second sample, which does not exclude individuals who have a spouse present, is also considered. In the lower half of Table 1, this sample is shown to contain 35,317 males and 33,570 females of whom 1,345 and 1,668 respectively are on social assistance, with further subsets of 715 and 771 individuals respectively who also work. In this larger sample, the proportion of individuals on social assistance is smaller because the presence of a spouse often signals improved circumstances through a second income. Nevertheless, the absolute number of individuals who work while on social assistance increases substantially for both genders. Table 1 indicates that, in both samples, the average number of hours worked per month conditional on labour force participation is about eighty, with males working slightly more hours than females and individuals with a spouse present working fewer hours than those of the same gender who had no spouse present. Table 1 also indicates substantial time dependence in social

[14] Gender determination is based on the LMAS variable SEX88. Data in Table 1 are unweighted.

[15] Social assistance receipt is determined by whether an individual received income from this source at any time during 1989 (LMAS variable Q138_89=1). Warburton (1998) notes that, when compared with administrative data for British Columbia to which as a government employee he has access, the LMAS appears to underreport the incidence of social assistance. It is conceivable that underreporting may be most severe by individuals most sensitive to program parameters, in which case our results would suggest more modest effects than would be the case in fact. However, this conjecture cannot be examined given available data.

[16] One is deemed to be in the labour force if one supplies positive hours in 1989 (LMAS variable DVTOHR89).

assistance status: Approximately 50% of the males who were on social assistance in 1989 were also in this program in the previous year; this percentage is, at about 70%, considerably higher for females. Individuals who were not on social assistance during 1989 were almost certainly not on social assistance during 1988. Time dependence is also noteworthy with respect to participation in the labour force: Between 83% and 86% of men who were on social assistance and worked during 1989 also worked in 1988 and these percentages are almost as high for females. Individuals who were not on social assistance in 1989 were even more likely to have been in the labour force during 1988. The degree of dependence in both processes underscores the need to work with data which can take this important force into account. In principle, detailed information on personal characteristics and adequate attention to heterogeneity might account for the underlying forces that lead to persistence. However, available data do not supply the necessary information on personal circumstances and including heterogeneity in the current context would considerably complicate the estimation framework. Accordingly, the individual's experience in the previous year is used as a fixed effect.[17]

A possible problem with the second sample in Table 1 is that eligibility for social assistance depends on family circumstances while the LMAS observations pertain to individuals and their own characteristics. However, the survey does include certain variables which describe the relation of the individual to the rest of the household and the involvement of family members in the labour force. The former includes whether the individual is the head of household.[18] The latter

[17] For a discussion of this specification and of heterogeneity see Nakamura and Nakamura (1985), page 275.

[18] In this paper, the dummy variable Head equals 1 when the LMAS variable RELHD89=1 and is zero otherwise.

includes a variable which reports the number of family members with full-time full-year jobs in 1989.[19] An additional conditioning variable is a spouse-present indicator.[20] These variables are used in the econometric analysis below.

The dependent variable in the Probit equation (1) above is the dummy variable described earlier. The hours variable in the H_j^i, $i = 0, 1$, equations 5 and 6 consists of hours-per-month (DVTOHR89/12). The matrix X in the Probit equation includes a constant, a wage rate,[21] the monthly social assistance guarantee level, the social assistance claw-back rate, whether the individual was on social assistance in 1988, his or her immigration status,[22] whether in 1989 the individual received pension income (LMAS Q137B89), his or her region of resi-

[19] This is variable FAMFYFT9 in LMAS. The survey also reports annual family income (FEARN89). However, this is reported in steps censored at 65 or $66,999. This amount is too low to be believable and it appears to be inconsistent with the individual's annual earnings which in this sample has a maximum of $166,880. Accordingly, this variable was not included in the analysis.

[20] This equals 1 when HHCHAR89=2 or 3 and is equal to zero otherwise.

[21] This is the LMAS variable DVWAGJ9 when it is positive and the unconditional predictions from a Heckman (1979) selection-adjusted wage equation when DVWAGJ9 is equal to zero. A number of other possibilities (e.g. unconditional predictions for all observations and conditional predictions for the unknown values) are clearly possible and these have been considered. In the main sample, 82% of men and 77% of women have known wages; using unconditional predictions for everyone would introduce forecast errors in the vast majority of cases. The preference for unconditional forecasts is discussed in CSS, footnote 28.

[22] Immigrants are individuals born outside Canada (LMAS COB=2).

dence,[23] age,[24] educational attainment,[25] whether the individual is a visible minority (LMAS VISMIN1=1), disabled,[26] married (LMAS MS89=1) has own children aged five or less (LMAS OWNKID189+OWNKID289) own children aged six to fifteen (LMAS OWNKID389), or other children under sixteen years of age (LMAS OTHKID189).[27] When individuals with a spouse present are included in the sample, the Probit equation includes, in addition, the three variables noted in the previous paragraph. The matrix Z in the Tobit equation excludes the dummy variable signifying receipt of pension income. The Tobit equation for individu-

[23] Five regions are distinguished. Atlantic (the four maritime provinces), Quebec, Ontario, Prairie (Manitoba and Saskatchewan) and the West (Alberta and British Columbia) - the latter being the omitted category.

[24] Four categories are distinguished. Less than 25, 25-34, 35-44 and over 44 years of age - the latter serving as the omitted category.

[25] The four levels distinguished consist of Incomplete High School, Completed High School, Trade Certificate and the omitted category of Post-Secondary Education.

[26] The dummy variable Disab. Limited below signifies a work-limiting disability (LMAS DISAB89 = 1). It should be noted that disability determination may be influenced by applications for disability status which may, in turn, depend on possible additional pecuniary benefits. Applications are approved only for individuals with a demonstrated need and the fact that, in this study, an individual is deemed to be disabled only if he/she is 'limited at work' suggests that the possibility of endogeneity is minimized. It should be noted that, while dropping the disability variable leads to clearly poorer fits, it does not affect the general conclusions reached in this study. A wider set of forces may be at work involving not only the parameters of the social assistance program but also those of other social programs, as well as various tax rates. Regional tax rates cannot be considered given the presence of regional effects in our equations and detailed information of the kind constructed in CSS is not available for other transfer programs.

[27] The LMAS children variables are integers with a maximum of four. In the tables, other children are denoted by 'O Ch'.

als off social assistance excludes, in addition, the program benefit and claw-back variables.[28] Both Tobit equations note the individual's labour force participation in the previous year and exclude the individual's social assistance status in 1988. When the spouse-present sample is used, the Tobit equations also include the spouse-present indicator, the headship variable and the number of jobs held by family members.

The construction of the parameters of the social assistance program used in this paper is a major undertaking. The National Council of Welfare (1990) provides a consistent profile of benefit rates, earnings exemptions and program tax rates for individuals and households by type and province, subject to a number of assumptions about the size of the municipality, the presence of two-tiered systems, the employability of the recipients, and the type of accommodation used by the recipient. The 1989 basic allowance reported in this publication includes consideration for food, clothing, shelter, utilities, and personal and household needs. Participants may also qualify for discretionary assistance for medical, dental and other special needs. The Council cautions that its estimates are maximum guarantees and that considerable discretion may be used by officials.

[28] In Moffitt's (1983) model, the individual is faced with two possible budget constraints, those on and off social assistance. If, taking possible stigma into account, the value of the indirect utility off social assistance is the highest, the individual does not participate and program parameters are not relevant to his or her labour-leisure choices; as in standard labour supply specifications, these are then influenced by the market wage rate. For a more detailed justification of this exclusion see Moffitt (1983). The wage and program claw-back rate are entered separately rather than in net wage form for two reasons: First, an implied constraint on coefficients is not imposed and, second, it is the pre-program wage rate which is relevant in the hours equation for non-participants.

The Council provides these estimates for individuals, lone parents with one child and couples with two children, by employability status.[29] Provincial regulations were used to extend the Council's basic allowance data to lone parents with two or more children and, in the second sample, to couples with other than two children. Thus the basic allowance in this study, which refers to benefits at zero hours of work, reflects variation by province, disability and the number of children.[30] For instance, the monthly, nominal,[31] allowance is $234 for employables in New Brunswick and $901 for lone parents in Ontario. The estimates for the additional support for children beyond the first one range from $36 per month in New Brunswick to $155 per month in Saskatchewan. In Quebec, the figure of $101 per month applies to one additional child only.

Provinces allow individuals on social assistance to retain a portion of earned income without reducing program support. The exemption formula varies by

[29] In this paper, disability and employability are treated as synonymous.While this is the only reasonable assumption given available information, it does gloss over possible differences in the treatment of lone parents with older children who are treated as unemployable in some but not in all provinces.

[30] The social assistance variables used in this study are similar to those in Christofides, Stengos and Swidinsky (1997a), Table 1. One difference in the application of the information in Table 1 is noteworthy. In the present paper, an individual is deemed to be disabled if he or she is reported to have a disability of any kind, while in the earlier paper that was only true of disabled individuals who are reported to be limited at work. This definitional difference is worth exploring given the considerable uncertainty concerning how disability status is determined by program administrators.

[31] In Christofides, Stengos and Swidinsky (1997a) we experimented with a real variable - deflated by a provincial housing cost index. The results were very similar to those obtained using the nominal variable.

province. New Brunswick exempts the first $200 of income earned by a lone parent but reduces its support dollar-for-dollar when income exceeds this amount. Alberta exempts 100% of earnings up to $115; 50% of earnings between $116 and $200; 25% of earnings between $201 and $300; and 10% of earnings over $300. The 1989 minimum and maximum marginal tax rates can be obtained from the National Council on Welfare (1990) and are also reported in Christofides, Stengos and Swidinsky (1997a), Table 1, columns 2, 4, 6 and 7. This table also reports the income level in 1989 dollars at which the maximum rates apply. For example, British Columbia exempts the first $50 earned by an employable individual and levies a 75% tax on income above this level, while Prince Edward Island levies a constant tax rate of 80% on all earnings. Some provinces calculate exemptions based on gross earnings, rather than earnings net of work-related costs. The rate structure is used to calculate the amount exempted given the individual's income. What, in short form, is defined in this study as the 'claw-back' rate, i.e. the proportion of earned income effectively lost because of decreased welfare benefits (1-exemption/earnings),[32] is an amalgam of two program parameters (the earnings exemption and the tax rate) and involves a number of assumptions about labour market experience (wages and hours). In CSS, the average rate constructed in this fashion was preferable to the marginal rates themselves because, in some cells, the latter are equivalent to provincial dummies. Thus, the claw-back rate which, given regional effects, is identified through the variation in wages and the non-linear transformation of marginal rates, should not be interpreted narrowly as a tax rate. Because a number of individuals do not work, their notional income is imputed

[32] This is constrained to be greater than or equal to the lowest marginal rate and it increases smoothly with earnings above the exemption.

using the individual-specific forecast wage and forty hours per month, or the approximate, unconditional, mean number of hours per month. Thus the social assistance claw-back rate used here varies by province, employability, household type and the individual-specific wage rate. In light of the interest that centers on this variable, results based on variants of this regressor are also reported. One alternative fixes the hours worked to eighty per month (the approximate conditional mean in Table 1), while another variant fixes the wage rate at the sample mean for those who work - $11.20 per hour. This latter variant reduces the claw-back variable to a function of program parameters only, namely the program tax rate and the allowable income exemption.[33]

5. Estimation Results

5.1. Main Results

Tables 2, 3 and 4 report separate equations for males and females. Each table contains results from both individually and system-estimated Probit and Tobit equations. The former are presented because they link the present study to earlier work on the component decisions. The system results are referred to as 'E/S Probit' and 'E/S Tobit' after endogenous selection. In addition to coefficients and standard errors, which are presented for both sets of results, columns 5 and 10, in Tables 2, 3 and 4 report the results of detailed scenarios which were considered using the system estimates. In Table 2, columns 5 and 10 relate to the probability of social assistance receipt. In Tables 3 and 4, columns 5 and 10 consist of two

[33] Thus the claw-back rate is exogenous if these program parameters can be treated as such. Warburton (1998) has suggested that this may not be so in British Columbia.

parts: The first part relates to the probability of participation in the labour force while the part after the semi-colon reports effects on expected hours. While most attention is given to the system estimates, notable differences between the two sets of results will be discussed and connections to the existing literature noted.

Table 2 deals with participation in the social assistance program. The wage rate has a significant negative effect on the probability of participation[34] in social assistance for men and the stand-alone Probit equation for women. This effect was also noted in Charette and Meng (1994) and Christofides, Stengos and Swidinsky (1997a and 1997b). In a similar vein, Barrett (1996) reports that a higher minimum wage rate encourages exit from the social assistance state. The program variables in the equations for males are not significant for two-tailed tests at the 5% level. They are of greater interest in the equations for females. In the Binomial Probit equation, there is evidence that higher benefits increase the probability of participation in social assistance. However, the coefficient on this variable, although positive, is not significant in the context of endogenous selection. The claw-back variable has the effect that would be expected in the context of the static labour supply model: An increase in the claw-back rate reduces program participation and this effect is robust across Table 2, columns 6 to 9. These results are generally more consistent with those obtained for female heads by Charette

[34] In the Probit model, the expected value of the index P_j is given by $F(X\alpha)$ and the effect of a continuous variable k on the expected value is given by $f(X\alpha)\alpha_k$, where X is the mean vector, F is the cumulative and f the probability standard normal density function. Since F and f are positive, the qualitative effect of a variable on the expected value is given by the coefficient sign. As most of the variables in the model are not continuous, their quantitative effect on the probability of social assistance receipt is evaluated by calculating the expected value $F(X\alpha)$ under different scenarios.

and Meng (1994) than those in CSS. Their quantitative significance is, of course, a matter of considerable policy interest and we now turn to this issue.

The variables just considered are the only continuous variables in Table 2. Row 1 of columns 5 and 10 in Table 2, indicates the predicted probability of social assistance receipt for the 'base' group[35] that is suggested when all remaining variables assume the value of zero - and the continuous variables are at their means. Thus, the predicted probability for a Canadian-born male who has no pension income, resides in British Columbia, is over forty-four years of age and so on is 8.4% and this happens to be very similar to the fraction of positive observations for the sample as a whole; this is 8.8% and it appears at the bottom of the column. In the case of females, the predicted probability for the base group is, at 3.3%, considerably lower than that (15.8%) for the sample as a whole. Rows 2 to 4, in columns 5 and 10 of Table 2, which relate to the three continuous variables, show the effect on the base probability of allowing each of these variables in turn to increase to its maximum logarithmic value. This change is much larger than the 21.6% decrease in benefits implemented in Ontario in October 1995 and hence the continuous variables, including the program variables, are indeed given a chance to affect relevant outcomes. In the case of benefits, for example, the increase from the mean to the maximum value would correspond to the increase that would be involved if a single, employable, individual in Ontario were granted the benefits appropriate for an unemployable adult with five children. Such an increase in the wage rate would reduce the probability of social assistance receipt

[35] It should be noted that this base group is constructed so that the tables can be easily intepreted. It is not intended to be 'representative' and predicted probabilities for base groups need not be close to the overall sample ones.

to essentially zero for both genders (by -8.4% from 8.4% for males and by -2.6% from 3% for females). Thus the wage rate is seen to have a strong quantitative effect. The program benefit and claw-back rate which, in the case of females, are generally statistically significant with expected signs are seen to have negligible effects. The largest effect is that for the benefits variable: An increase to its maximum value would raise the probability of social assistance receipt by 1.3% points. However, this also the only program variable which is not significant in the equations for females.[36]

Table 2 suggests that only male immigrants are less likely to be on social assistance; this variable is associated with a 4% point decline in the probability of social assistance receipt to 4.4% - relative to the 8.4% for the base group. Charette and Meng (1994) report positive but insignificant coefficients for female

[36] In Ontario, the social assistance caseload declined from just over a million beneficiaries in June 1995 to 644,300 beneficiaries by February 1999. This decline of some 40% has often been associated in th epopular press with the reform of the social assistance system carried out by the Harris government after it assumed power in June 1995. One has to be very cautious about jumping from the cross-sectional results above to the effects over time of the changed benefit, administrative and legal landscape that was brought about by the Harris government. Certainly for men, it is difficult to detect any program effects in the results above and so the argument must be that the decline in the caseload for men (the Ontario government data at www.gov.on.ca/CSS/page/ are not broken down by gender) is due to the changes in the administrative and legal context and the improving labour marklet conditions - on the last possibility, see the regional effects below and note, also, that the decline in the caseload begun prior to the Harris reforms as the economy begun to recover. The effects for women are better-established but nevertheless weak so the argument is essentially the same as that for men. Below, we note the very substantial time dependence in social assistance participation which suggest that long-run effects will be substantially larger than the impact effects in Table 2.

heads. Abstracting from their cohort and assimilation effects noted earlier, Baker and Benjamin (1995) note that immigrant families are less likely to be on social assistance. The differential effect for men and women may arise because immigrant screening for men may be more focussed on labour market skills and result in better job matches than is the case for women. The visible minority indicator is not a useful predictor of the probability of social assistance receipt.

Pension income is associated with a lower probability of social assistance receipt regardless of estimation method and gender. The associated quantitative effects, in columns 5 and 10 of Table 2, are -4.5% for males and -2% for females.

The regional dummies are intended to capture differences in social assistance programs[37] and possible differences in labour market tightness across the regions of Canada. On the whole, the Atlantic provinces and Quebec behave much like the omitted category (West), while social assistance receipt in Ontario and the Prairie provinces is less likely. In the case of females, these regional effects are at most -2% for the Prairies. In the case of males, the *gap* between the province with the lowest likelihood of social assistance reliance (Ontario) and that with the highest (Quebec) is 8.2% (4.7% minus -3.5% in column 5, Table 2). This gap, while consistent with differences in labour market tightness, may reflect more than that (e.g. administrative practices) given the estimated effect for the Atlantic provinces where unemployment is even higher.

Coefficients on age variables are significant for males only and then only for the first two categories. They suggest that younger men are less likely to be on social assistance than is the case for those over 44 years of age by approximately 5%

[37] These might include program features not captured by the two social assistance variables as well as differences in the administration of these programs.

points. Relative to individuals with post-secondary education, less well-educated men and women are more likely to receive social assistance. This effect is especially strong for the best-educated female group where the quantitative impact is as high as 8% points.

Work-limiting disability is associated with a statistically significant and very substantial increase in the probability of social assistance receipt. This holds for both males (12.5%) and females (8%).

Married individuals have a lower probability of social assistance receipt which ranges from -2.7% points for males to -1% points for females. This may be due to income support from a spouse, though it is noteworthy that most coefficients are not significant at the 5% level. The presence of own children, particularly young ones, increases the probability of social assistance receipt. It is noteworthy that while these variables are statistically better established for females, the quantitative effects are stronger for males. Thus, the probability of social assistance receipt for a male in the base group who, in addition, has a child under five years of age increases by 9.5% points while that for a female increases by 6% points.

It is worth emphasizing that the results above were obtained in the context of a specification which takes into account the individual's social assistance experience during 1988. The extent of persistence was first encountered in Table 1 and it is very clear in all the equations of Table 2. The quantitative effects are very large for both genders: The probability of social assistance receipt by a male in the base group jumps from 8.4% to 54.4% if this individual also received assistance in the previous year. The strength of this variable suggests that a number of personal characteristics which are not reported in the LMAS and other similar data sets are critically important and must, at the very least, be proxied as in

Table 2. It is also worth noting that, if beyond individual effects, there is also genuine time-dependence, the effects of shocks and changes in variables will be felt over time.

Table 3 reports Tobit hours-of-work equations for males estimated both in stand-alone form (Table 3, columns 1,2,6 and 7) and as a system with the Probit equations reported in Table 2 (Table 3, columns 3,4,8 and 9). Separate equations for individuals on and off social assistance are reported. Columns 5 and 10, Table 3, report the predicted probability of labour force participation (32% and 34% for men on and off social assistance respectively) and expected hours per month (12.6 and 18.7 for men on and off social assistance respectively) for base groups which, as noted earlier, are constructed for the reader's convenience and will not be representative. Microeconomic theory suggests that the wage rate effects on effort are ambiguous. In Table 3, the estimated coefficients suggest an upward sloping function for males off social assistance and a backward bending supply function for males on social assistance. Though statistically significant, the former is quantitatively weak given that, as the wage rate increases from its mean to its maximum value, the probability of labour force participation increases by 2% points while expected hours increase by 2.3 per month.[38] The effects while

[38] In the case of Tobit, the derivative of the unconditional mean of H_j^i with respect to a continuous variable k is given by $F(Z^i\beta^i/\sigma^i)\beta_k^i$, $i = 0, 1$, where Z^i is the mean vector. Since most variables are categorical, the implied quantitative effects are assessed by calculating, under different assumed conditions, the probability $F(\cdot)$ of a non-limit observation, i.e. the probability of labour force participation, as well as the unconditional expectation of hours per month which is given by $E(H^i) = Z^i\beta^i F(Z^i\beta^i/\sigma^i) + \sigma^i f(Z^i\beta^i/\sigma^i)$.These effects are reported in columns 5 and 10, Tables 3 and 4. Effects on expected hours conditional on no censoring can be read from the estimated coefficients in Tables 3 and 4. Note that, here, conditioning refers to censoring

on social assistance are much stronger: This increase in the wage rate would reduce the probability of labour force participation to 1% (32% base plus an effect of -31%) and expected hours to 1 per month (12.6 base plus -11.6). Since a higher wage rate means that less hours are required to reach the program earnings exemption and claw-back rates beyond the exemption are often 100%, a higher wage rate is likely to reduce effort. This result is not surprising and was also found by CSS for single individuals - the wage rate effects on hours supplied by the lone parents in the CSS sample were are not significant.

An examination of the social assistance program variables suggests that higher program benefits reduce effort. These effects appear quite sizeable: When the benefit level increases from its mean to its maximum value,[39] the probability of labour force participation declines by 14% points and expected hours by 6.4 per month. The social assistance claw-back variable does not have a well-established coefficient in Table 3.

Immigrant men who are off social assistance work significantly less hours per month, an effect which is not present for immigrant men on social assistance. Visible minority men who are off social assistance work significantly longer hours than the control group. These effects and those on the implied probability of labour force participation are quite substantial - Table 3, columns 5 and 10.

There are no significant differences in labour force involvement between Ontario and the West but this is generally so for other parts of the country. For instance, males on social assistance in Atlantic Canada have a probability of labour

and not explanatory variables - the latter is not shown.

[39] It should be noted that this increase is determined by provincial regulations and is modest - see Table 1 in CSS.

force participation which is 26% points lower than in the West. In addition, their expected hours of work per month are 11.1 lower than in the West. Effects of similar orders of magnitude hold for Quebec and these are present, albeit in muted form, for men off social assistance. If these are manifestations of labour market tightness, they would justify not imposing the standard labour-leisure structure on the data.

On the whole, older and less educated men have a more marginal attachment to the labour force, both in terms of participation and hours, than individuals in the omitted categories. Some of these effects, e.g. those for the youngest group on social assistance, are comparable in size to the strongest regional effects. In the case of men on social assistance, the disability variable, though carrying a negative sign, is not significant. This may be the result of an insufficient number of individuals in this cell. Disabled individuals off social assistance are 9% points less likely to be in the labour force and work 6 fewer hours per month than the control group.

The coefficients on marital status are positive for men who are on and negative for men who are off social assistance, though significance is only attained in the stand-alone Tobit equations. This may reflect need - note that own children aged 6-15 also increase the labour market involvement of men on social assistance.

Table 4 contains results for women. As there are a number of similarities with the results for males, particular attention is paid to the notable differences between the results in Table 3 and 4. The probability of labour force participation and expected hours of work for the benchmark groups appear, with the same caveat as for men, in row 1, columns 5 and 10, Table 4.

In contrast to the results for men off social assistance, the coefficient on the

wage rate for women who are off social assistance is negative and significant. This is not an unusual finding.[40] The reduction in the labour force participation rate and hours of work as the wage rate increases from its mean to its maximum value are 20% points and 9.6 hours respectively - column 10, Table 4. As for men, the social assistance variables continue to have negative coefficients in the equations for women on social assistance and the quantitative effects for the benefits variable are not negligible. However, neither variable is statistically significant.

The coefficients on the regional dummies in Table 4 suggest that, as in the case of men, labour market involvement by women is significantly lower in the Atlantic Provinces and Quebec than is the case in the West - the quantitative effects for these regions in column 5, Table 4, are especially substantial. For women on social assistance, it is also lower in Ontario and the Prairies, though Prairie women who are off social assistance are more likely to participate in the labour force and to work longer hours than those in the West.

The sign pattern for the age and education variables is broadly similar to that for men. Work-limiting disability has the anticipated negative coefficients and these are generally statistically significant. The quantitative effects for women on social assistance are weaker than those for women off social assistance, a pattern also present in Table 3, presumably because there is more variance in the labour market involvement of individuals off social assistance. The hours equations for women on social assistance do not afford any role to the marital status and number of children variables. For women off social assistance, there is strong evidence of negative supply effects arising from the number of children virtually regardless of

[40] For a review of findings on hours equations which includes those in Nakamura and Nakamura (1981) for Canada, see Killingsworth (1983).

age.

The correlation coefficients between the error term in the Probit and each of the Tobit equations in the endogenous selection specification are always positive and significant at the 10% level in all but the case of hours supplied by men on social assistance. This finding, which is a feature of the system approach adopted in this paper, appears robust to changes in specification. That is, unobservables which tend to increase the probability of being on social assistance also increase labour market involvement, especially for men and women who are not already on social assistance. This finding may hint of forces pertaining to financial hardship and the need and propensity to work which are not adequately captured by the variables in Tables 2-4. This interpretation would be consistent with the general tone of the findings in this paper which tend to accord only a limited role to program variables.

As in the social assistance participation equation, the individual's 1988 experience may serve as a kind of fixed effect which captures important unobserved characteristics. Thus, Tables 3 and 4 also condition on labour force participation in 1988 and this variable is seen to be extremely powerful. Labour force participation during 1988 increases the probability of participation in 1989 by 40 to 60% points. Expected hours of work also increase dramatically by amounts which are especially large for individuals off social assistance (e.g. 116.8 hours per month for men). These effects are reasonable given the data in Table 1. Experience during 1988, which is not included in the benchmark calculations of row 1, columns 5 and 10, Tables 3 and 4, helps bring the figures in those cells to more representative levels given that most individuals, but especially those off social assistance, were in the labour force during 1988 - see Table 1.

The system approach to modelling social assistance and labour supply has clear conceptual and econometric advantages. It is also noteworthy that it often implies different quantitative responses. For instance, the coefficient on the social assistance benefit variable for females in Table 2 is considerably smaller in the system-based Probit equations than in the stand-alone Probit equation. This may be because the system approach takes into account the endogenous hours response. For instance, in Table 4, an increase in the benefit variable will induce a considerably smaller decline in labour supply (and hence increase in social assistance participation) in the system case than in the stand-alone Tobit case.

5.1.1. Sensitivity Analysis

As noted in section 4, the construction of the social assistance program variables is a complex undertaking which requires many steps and assumptions. The benefit variable used in this study is based on very detailed information concerning the individual's place of residence, disability and size of family. It is difficult to see how this variable might be improved. The claw-back variable incorporates details of program earnings exemptions and tax rates which also vary by province, disability and family size. These, too, are difficult to improve upon. However, the claw-back variable also includes assumptions about the number of hours worked and the predicted wage rate. Two alternative variables, one based on 80 hours per month and another which sets the wage rate for all individuals at its average level conditional on participation ($11.20) were also considered; in the latter, the claw-back variable depends solely on program parameters.

The results on the assumed 80 hours per week specification indicate that the general sign pattern, coefficient significance and, in most cases, coefficient size

remain unaffected. Moving to the constant wage rate specification and comparing to the 80 hours per month and constructed wage rate benchmark of Tables 2 to 4, it is clear that what sometimes appears as a claw-back effect may actually be a wage rate effect: When the constant-wage claw-back variant is used it is not significant and, where it previously was negative and significant, the wage rate now has a stronger, negative, coefficient. These results are not presented in detail.

5.2. The Second Sample

Tables 5 and 6 present partial results based on the larger sample which does not exclude individuals who have a spouse present. All equations include a spouse-present and a headship indicator as well as the number of family members with a full-year full-time job. The spouse-present variable reduces the probability of social assistance receipt in all equations (by -9.1% and -1.8% for males and females respectively) but it is not significant in the system Probit equations. Household Heads are less likely to be on social assistance if they are male and more so if they are female. This may arise because female Heads may reside in households with elderly parents or relatives who have no income. These effects, in columns 5 and 10, Table 5, are -9.5% and 1.8% points respectively. Additional family jobs depress the probability of social assistance receipt for both genders: A single family job would reduce this probability by 12.3% points for males and 5.4% points for females; the maximum for this variable is five jobs. The inclusion and performance of these controls suggests that this larger sample may yield useful information on the performance of the program variables.

The switch to this larger sample appears to have generally strengthened the benefit variable which is now positive in all equations and is significant for females.

The size of this coefficient is now nearly double that in the sample of Table 2 but, despite this, the effect in column 10, Table 5, is only a 3.2% point increase in the probability of social assistance receipt as the benefit variable increases from its mean to its maximum.[41] The claw-back variable, which as seen above, may actually capture a wage effect is no longer significant. These results are still based on specifications which control for all the variables in Table 2 and, in general, these variables have the sign patterns and coefficient sizes reported earlier.

In the interests of brevity and because this is where the social assistance variables appear, Table 6 reports hours equations only for individuals on social assistance. The benefit variable has the negative sign seen in Tables 3 and 4 but it is now statistically weaker. By contrast the tax variable is consistently negative and significant, though this may be a wage effect. The wage rate itself has negative, significant, coefficients in the Tobit equations for women but it is not significant in the equations for men.

6. Conclusions

The labour force and social assistance participation decisions are clearly interrelated for with sufficient hours of work at a high-enough wage rate an individual would not qualify for support. This paper explores this interaction and attempts

[41] When, for males, the sample is restricted further to Heads only, the benefits variable is positive and significant in both the Binomial Probit (coefficient of 0.24 with a c/se of 2.19) and E/S Probit equation (coefficient of 0.02 with a c/se of 3.69). The claw-back variable is only significant in the E/S Probit equation (coefficient of -0.003 with a c/se of 2.34). For female Heads, the sample analysed by Charette and Meng (1994), the program variables perform as in Table 5, albeit more strongly.

to reconcile what on the surface of it appear to be the different results obtained when this approach is used and when the social assistance participation decision is studied independently of the labour market decision. Tobit equations for hours worked by individuals on and off social assistance are estimated using Full Information Maximum Likelihood, allowing for endogenous selection of the individual's social assistance status through a simultaneously estimated Probit equation. In addition, stand-alone Tobit and Probit equations are reported both to link up with the earlier Canadian literature and to examine the extent to which results obtained are sensitive to the econometric methodology adopted. Separate samples of males and females who do not have spouses are considered. In addition, a second set of samples for males and females who have a spouse present is considered both because these samples are larger and because they correspond more closely to some of the samples studied in earlier literature. Both sets of samples may involve individuals who have children present. An important feature of the 1988-89 Longitudinal LMAS data used is that it is possible to condition on the past labour force and social assistance participation by these individuals. This makes it possible to take into account the very considerable social assistance and labour force participation persistence that is evident in the raw data of Table 1.

The fit of the various equations is reasonable and many personal characteristics such as age, education, marital status, the age and number of children and disability are, in general, important determinants of these decisions. Of particular relevance to this study is the role of the social assistance variables. Looking first at the social assistance participation equations and the case of males, the benefit level, which is not useful in the sample of males without a spouse present, acquires a positive coefficient and becomes significant at lower size levels in larger samples

where individuals with a spouse present are included, particularly in a sample of men who are heads of household. However, the quantitative contribution of this variable is very small. For males, the claw-back rate (the proportion of income lost because of decreased welfare benefits as earned income rises) has nothing useful to contribute to the social assistance participation equation.

In the case of females, the benefit variable always carries a positive sign and is significant, particularly in larger samples. The claw-back variable has the anticipated negative coefficient and this is significant even in the system-based Probit equation. However, this result is sensitive to the construction method for this variable and it may simply be a wage effect. That is, the higher the wage rate the lower the propensity to rely on social assistance and once this ingredient of the claw-back variable is removed, leaving it a function of the underlying program parameters only, this variable is no longer significant. In addition, the claw-back variable is not significant in the larger sample of females who have a spouse present.

The results obtained for females reconcile previous findings, most of which deal with the case of females: Charette and Meng (1994), for example, treat the social assistance participation decision as independent of labour supply, work with an LMAS sample of female heads of household, and find that the benefit level has a significant positive coefficient in social assistance participation equations. By contrast, CSS treat the social assistance and labour supply decisions as interdependent, work with smaller LMAS samples of single females or lone mothers and find that the benefit variable is not useful. The results in this paper suggest that the estimation method does make a difference in the smaller sample of females who have no spouse present: When the interdependence of social assistance and

labour supply is taken into account and a system estimation method used, the benefit variable, which is significant in the stand-alone Probit equation, is no longer significant. However, in the larger sample of females with a spouse present and particularly those who are family heads the benefit variable is always significant. Thus, the difference in earlier findings for females is resolved as a sample size, rather than an estimation method, issue. It should be noted that in spite of this reconciliation of earlier findings, the quantitative effects of the benefit variable is modest. The performance of the claw-back variable in Charette and Meng (1994) is consistent with the mixed results obtained for this variable here.

In the system-based hours equations for males and females on social assistance, the two program variables always have negative signs. In the smaller samples, from which individuals with a spouse present are excluded, the benefits variable is significant at the 5% level for males and has a noteworthy quantitative effect on hours; however, the statistical significance of this variable is not maintained in the larger sample. The benefits variable is never significant in the Tobit equations for females. In the larger sample, where individuals with a spouse present are included, the claw-back variable is consistently significant; it should be noted, however, that this variable is never significant in the smaller sample. Thus, program variables do not appear to have reliable effects on labour force involvement.

There is, then, some evidence that program benefits affect participation in social assistance particularly by females. The size of this effect is smaller in the system than in the stand-alone participation equations in earlier literature; maximal impacts are in the order of 2 to 3% point increases in the probability of social assistance receipt as benefits rise from their mean to their maximum value. The labour supply effects of the benefit variable are less well established.

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Table 1

Sample Characteristics

	Males	**Females**
First Sample: No Spouse	*6932*	*7169*
On Social Assistance	*611*	*1135*
In the Labour Force	305	558
Hours per month	87.0	81.7
On SA in prev. year (cell %)	50.8	71.3
In LF in prev. year (cell %)	84.9	84.1
Not On Social Assistance	*6321*	*6034*
In the Labour Force	5370	4971
Hours per month	149.1	140.8
On SA in prev. year (cell %)	3.4	4.4
In LF in prev. year (cell %)	97.0	96.8
Second Sample: Spouse Present	*35317*	*33570*
On Social Assistance	*1345*	*1668*
In the Labour Force	715	771
Hours per month	81.0	77.9
On SA in prev. year (cell %)	48.3	67.2
In LF in prev. year (cell %)	86.0	82.7
Not On Social Assistance	*33972*	*31902*
In the Labour Force	27843	23521
Hours per month	153.1	120.4
On SA in prev. year (cell %)	1.9	2.0
In LF in prev. year (cell %)	97.0	94.9

Table 2

Participation in Social Assistance by Gender

	Males					**Females**				
	Bino. Probit		E/S Probit			Bino. Probit		E/S Probit		
	coef	c/se	coef	c/se	effect	coef	c/se	coef	c/se	effect
Constant	-1.23	1.81	-1.11	1.88	0.084[a]	-3.48	5.48	-2.77	4.34	0.030[a]
ln Wage	-0.04	4.93	-.22E-1	2.62	-0.084[b]	-.18E-1	2.40	-.10E-1	1.35	-0.026[b]
ln SA Ben	0.01	0.06	.12E-2	0.01	0.000[b]	0.25	2.51	.1548	1.53	0.013[b]
ln SA ClBk	0.01	0.39	.26E-1	1.57	0.001[b]	-.53E-1	2.58	-.55E-1	2.61	-0.002[b]
SA in 1988	1.91	29.96	1.39	22.61	0.460[c]	1.94	36.00	1.73	32.45	0.410[c]
Immigrant	-0.46	3.07	-0.33	1.98	-0.040[c]	-0.08	0.76	-0.12	1.03	-0.010[c]
Pension Inc	-0.47	3.86	-0.38	3.13	-0.045[c]	-0.68	7.01	-0.68	7.00	-0.020[c]
Reg: Atl	-0.09	1.13	-0.06	0.93	-0.009[c]	-0.07	0.94	-0.09	1.24	-0.010[c]
Reg: Que	0.13	1.67	0.26	2.85	0.047[c]	0.04	0.54	0.07	0.84	0.010[c]
Reg: Ont	-0.35	3.42	-0.27	2.81	-0.035[c]	-0.21	2.66	-0.15	1.64	-0.010[c]
Reg: Pr	-0.22	2.37	-0.01	0.12	-0.002[c]	-0.30	3.54	-0.30	3.58	-0.020[c]
Age: $\leq$ 24	-0.45	4.90	-0.52	5.48	-0.055[c]	0.14	1.70	0.02	0.26	0.000[c]
Age: 25-34	-0.35	4.30	-0.39	4.52	-0.046[c]	0.02	0.23	-0.09	1.12	-0.010[c]
Age: 35-44	-0.04	0.46	-0.14	1.63	-0.020[c]	0.08	0.95	-0.03	0.30	0.000[c]
Ed: <HSch	0.28	3.71	0.22	3.00	0.039[c]	0.64	9.63	0.67	9.80	0.080[c]
Ed: HSch	0.16	1.89	0.15	1.80	0.025[c]	0.25	3.40	0.26	3.45	0.020[c]
Ed: TrCert	0.13	1.00	0.05	0.35	0.008[c]	0.32	2.74	0.33	2.45	0.030[c]
Vis Minor	-0.17	0.76	-0.17	0.67	-0.023[c]	0.20	1.32	0.26	1.39	0.020[c]
Disab Lim	0.56	6.92	0.57	6.99	0.125[c]	0.64	9.23	0.64	9.02	0.080[c]
Married	-0.55	3.55	-0.20	0.98	-0.027[c]	-0.21	1.71	-0.20	1.76	-0.010[c]
Chld $\leq$5	0.50	2.42	0.46	2.51	0.095[d]	0.49	7.51	0.52	8.71	0.060[d]
Chld 6-15	0.11	1.33	0.16	1.82	0.025[d]	0.12	2.67	0.14	3.23	0.010[d]
O Ch $\leq$15	0.04	0.23	0.08	0.30	0.013[d]	-0.01	0.05	0.01	0.09	0.000[d]
N; Pve; %	6932	611	6932	611	0.088	7169	1135	7169	1135	0.158
Log Likel	-1245		-34656			-1554		-33195		
R^2	0.39					0.53				
Corr Pred	0.94					0.92				

Notes: (*a*) Predicted probability of SA receipt with continuous variables at their means and discrete variables at zero. (*b*) Effect on the probability of SA receipt of increase in continuous variable from its mean to its maximum value. (*c*) Effect on the probability of SA receipt of increase in dummy from 0 to 1. (*d*) Effect on the probability of SA receipt of increase in categorical variable from 0 to1 child. Columns 2, 4, 7 and 9, which report values for coefficients/errors, are in absolute terms.

Table 3

Supply of Hours by Males

	On Social Assistance					No Social Assistance				
	Tobit		E/S Tobit			Tobit		E/S Tobit		
	coef	c/se	coef	c/se	effect	coef	c/se	coef	c/se	effect
Constant	210.1	2.59	193.1	2.19	.32; 12.6[a]	-38.8	6.33	-43.9	7.14	.34; 18.7[a]
ln Wage	-3.4	3.95	-3.5	4.13	-.31; -11.6[b]	0.2	0.86	0.7	3.05	.02; 2.3[b]
ln SA Ben	-34.7	2.63	-29.9	2.12	-.14; -6.4[b]	n/a	n/a	n/a	n/a	n/a; n/a
ln SA ClBk	-1.5	0.68	-1.3	0.58	.00; -0.2[b]	n/a	n/a	n/a	n/a	n/a; n/a
LF in 1988	110.2	14.42	113.7	14.17	.60; 75.0[c]	181.0	44.97	168.8	48.78	.60; 116.8[c]
Immigrant	-28.4	1.31	-26.5	1.29	-.14; -6.6[c]	-13.6	3.30	-11.1	2.45	-.05; -3.4[c]
Reg: Atl	-70.2	7.96	-66.1	7.13	-.26; -11.1[c]	-18.7	6.03	-17.4	5.30	-.08; -5.1[c]
Reg: Que	-52.3	5.76	-47.8	5.00	-.22; -9.3[c]	-14.8	4.47	-15.4	4.20	-.07; -4.6[c]
Reg: Ont	16.7	1.46	17.1	1.32	.11; 6.4[c]	-4.7	1.52	-0.5	0.16	-.01; -0.1[c]
Reg: Pr	-61.7	5.41	-53.8	3.45	-.23; -10.6[c]	-3.4	1.06	1.1	0.31	.00; 0.5[c]
Age: $\leq$ 24	43.3	3.98	31.0	2.63	.20; 13.0[c]	16.7	4.61	22.1	6.06	.10; 8.7[c]
Age: 25-34	35.3	3.67	24.3	2.14	.16; 9.6[c]	17.0	5.10	19.9	5.81	.09; 7.7[c]
Age: 35-44	38.1	4.03	29.3	2.71	.19; 11.4[c]	13.2	3.58	12.3	3.33	.05; 4.6[c]
Ed: <HSch	-10.4	1.20	-12.3	1.24	-.07; -3.5[c]	-22.5	8.15	-27.9	9.25	-.11; -7.7[c]
Ed: HSch	9.9	0.99	9.4	0.80	.06; 3.2[c]	-6.5	2.32	-7.8	2.58	-.03; -2.4[c]
Ed: TrCert	31.2	1.77	13.5	0.56	.09; 4.8[c]	-2.6	0.60	-1.6	0.37	-.01; -0.4[c]
Vis Minor	-32.9	0.82	0.2	0.00	.00; 0.0[c]	24.3	4.18	26.9	4.30	.12; 10.8[c]
Disab Lim	-3.1	0.36	-8.6	0.93	-.05; -2.5[c]	-13.9	3.18	-20.9	5.65	-.09; -6.0[c]
Married	66.0	3.10	31.2	1.18	.20; 10.1[c]	-8.3	1.99	-3.4	0.86	-.02; -1.0[c]
Chld $\leq$5	-41.7	1.83	-16.6	0.55	-.09; -4.6[d]	28.7	2.77	18.2	1.36	.08; 7.0[d]
Chld 6-15	32.6	4.11	29.3	3.54	.19; 12.1[d]	-0.7	0.21	-0.7	0.20	-.01; -0.1[d]
O Ch $\leq$15	-1.9	0.09	4.1	0.11	.02; 1.4[d]	-2.3	0.42	-2.4	0.45	-.01; -0.7[d]
σ	60.2	23.54	60.2	24.42		80.0	101.22	84.1	162.96	
ρ	n/a	n/a	-0.05	0.45		n/a	n/a	0.88	50.33	
Log Likel	-1781		-34656			-31794		-34656		
Obs	611		611			6321		6321		
Pos Obs	305		305			5370		5370		

Notes: (*a*) Predicted probability of labour force participation F and expected hours E(H), in F; E(H) format, with continuous variables at their means and discrete variables at zero. (*b*) Effects on F; E(H) of increase in continuous variable from its mean to its maximum value. (*c*) Effects on F; E(H) of increase in dummy from 0 to 1. (*d*) Effects on F; E(H) of increase in categorical variable from 0 to 1 child. Columns 2, 4, 7 and 9, which report values for coefficients/standard errors, are in absolute terms.

Table 4

Supply of Hours by Females

	On Social Assistance					No Social Assistance				
	Tobit		E/S Tobit			Tobit		E/S Tobit		
	coef	c/se	coef	c/se	effect	coef	c/se	coef	c/se	effect
Constant	236.6	3.04	185.1	2.06	.46; 24.0[a]	-26.1	5.28	-23.3	4.82	.33; 14.2[a]
ln Wage	-10.9	11.52	-11.0	11.6	-.42; -18.1[b]	-0.8	4.04	-0.6	2.68	-.20; -9.6[b]
ln SA Ben	-20.8	1.73	-13.6	0.97	-.07; -4.9[b]	n/a	n/a	n/a	n/a	n/a; n/a
ln SA ClBk	-1.87	1.28	-2.3	1.27	-.02; -0.8[b]	n/a	n/a	n/a	n/a	n/a; n/a
LF in 1988	86.0	13.7	83.7	11.5	.40; 56.8[c]	163.8	50.33	157.1	57.30	.64; 113.5[c]
Immigrant	-14.6	1.06	-14.6	1.12	-.09; -6.1[c]	-1.5	0.42	-1.0	0.27	-.01; -0.3[c]
Reg: Atl	-57.3	8.07	-58.0	7.42	-.29; -17.6[c]	-8.3	3.09	-7.9	2.89	-.05; -2.3[c]
Reg: Que	-55.6	6.25	-56.0	5.53	-.28; -17.3[c]	-11.4	4.02	-11.0	3.50	-.06; -3.2[c]
Reg: Ont	-33.9	3.86	-37.2	3.99	-.20; -13.2[c]	2.2	0.87	4.2	1.49	.02; 1.5[c]
Reg: Pr	-29.8	3.57	-32.1	2.95	-.18; -11.8[c]	8.8	3.22	11.6	4.05	.06; 4.2[c]
Age: $\leq$ 24	8.5	.83	5.8	0.49	.03; 2.7[c]	19.2	6.88	16.7	6.20	.09; 6.3[c]
Age: 25-34	27.5	2.78	23.6	2.05	.13; 12.4[c]	27.1	10.60	24.7	9.11	.14; 9.8[c]
Age: 35-44	24.0	2.36	21.6	1.73	.12; 11.2[c]	26.7	9.49	25.2	8.55	.14; 10.1[c]
Ed: <HSch	-38.7	5.70	-36.3	4.63	-.20; -12.9[c]	-21.3	8.26	-26.9	10.07	-.13; -6.9[c]
Ed: HSch	-24.8	3.11	-24.3	2.80	-.14; -9.4[c]	-5.7	2.50	-8.0	3.30	-.05; -2.4[c]
Ed: TrCert	0.5	0.04	2.9	0.22	.01; 1.4[c]	-5.3	1.37	-6.6	1.48	-.04; -2.0[c]
Vis Minor	0.9	0.05	-1.3	0.08	-.02; -0.6[c]	5.7	1.09	5.2	0.98	.03; 1.8[c]
Disab Lim	-17.6	2.75	-14.5	1.93	-.09; -6.0[c]	-34.5	9.76	-39.0	12.6	-.18; -9.0[c]
Married	21.7	1.24	15.3	0.92	.08; 7.6[c]	-10.0	2.52	-7.7	1.88	-.05; -2.3[c]
Chld $\leq$5	0.3	0.05	2.5	0.42	.01; 1.2[d]	-10.3	3.09	-18.5	5.59	-.10; -5.1[d]
Chld 6-15	1.3	0.33	1.4	0.30	.00; 0.6[d]	-1.7	1.05	-5.5	3.08	-.03; -1.6[d]
O Ch $\leq$15	-13.9	0.99	-12.5	0.69	-.07; -5.3[d]	-17.1	4.80	-16.6	5.06	-.09; -4.6[d]
σ	69.5	31.8	69.4	37.9		65.2	97.65	66.9	159.97	
ρ	n/a	n/a	0.13	1.68		n/a	n/a	0.58	15.97	
Log Likel	-3343		-33195			-31794		-33195		
Obs	1135		1135			6034		6034		
Pos Obs	558		558			4971		4971		

Notes: (*a*) Predicted probability of labour force participation F and expected hours E(H), in F; E(H) format, with continuous variables at their means and discrete variables at zero. (*b*) Effects on F; E(H) of increase in continuous variable from its mean to its maximum value. (*c*) Effects on F; E(H) of increase in dummy from 0 to 1. (*d*) Effects on F; E(H) of increase in categorical variable from 0 to 1 child. Columns 2, 4, 7 and 9, which report values for coefficients/standard errors, are in absolute terms.

Table 5

Participation in Social Assistance by Gender (Spouse Present)

	Males					**Females**				
	Bino. Probit		E/S Probit			Bino. Probit		E/S Probit		
	coef	c/se	coef	c/se	effect	coef	c/se	coef	c/se	effect
Constant	-1.67	2.06	-1.48	3.04	0.150[a]	-4.29	8.02	-2.88	6.57	0.081[a]
ln Wage	-0.05	8.61	-0.03	6.97	-0.150[b]	-0.04	5.72	-0.03	4.66	-0.081[b]
ln SA Ben	0.15	1.72	0.12	1.63	0.021[b]	0.46	5.60	0.26	3.86	0.032[b]
ln SA ClBk	-0.02	1.13	-0.01	0.77	-0.001[b]	-0.01	0.64	-0.02	1.43	-0.002[b]
Sp Present	-0.25	3.21	-0.08	1.14	-0.091[c]	-0.27	3.09	-0.13	1.80	-0.018[c]
Head	-0.13	2.39	-0.12	2.44	-0.095[c]	0.24	4.12	0.11	2.40	0.018[c]
# Fam Jobs	-0.62	18.81	-0.44	13.16	-0.123[d]	-0.61	17.21	-0.52	15.08	-0.054[d]
N; Pve; %	35317	1345	35317	1345	0.038	33570	1668	33570	1668	0.050
Log Likel	-3161		-168627			-2856		-141054		
R^2	0.37					0.52				
Corr Pred	0.97					0.97				

Notes: (*a*) Predicted probability of SA receipt with continuous variables at their means and discrete variables at zero. (*b*) Effect on the probability of SA receipt of increase in continuous variable from its mean to its maximum value. (*c*) Effect on the probability of SA receipt of increase in dummy from 0 to 1. (*d*) Effect on the probability of SA receipt of increase in categorical variable from 0 to1 child. Columns 2, 4, 7 and 9, which report values for coefficients/errors, are in absolute terms. All equations were estimated with the same specification as in Table 2 with the additional regressors, shown above, to control for family circumstances.

Table 6

Supply of Hours by Individuals on Social Assistance (Spouse Present)

	Males					**Females**				
	Tobit		E/S Tobit			Tobit		E/S Tobit		
	coef	c/se	coef	c/se	effect	coef	c/se	coef	c/se	effect
Constant	-117.2	1.70	-17.4	0.22	.25; 10.0[a]	52.9	0.79	109.0	1.52	.36; 17.2[a]
ln Wage	0.6	0.86	1.2	1.34	.23; 15.8[b]	-9.6	9.03	-9.4	14.01	-.36; -17.2[b]
ln SA Ben	4.9	0.46	-6.5	0.54	-.03; -1.0[b]	0.9	0.08	-9.7	0.89	-.04; -2.5[b]
ln SA ClBk	-4.5	3.34	-4.5	3.33	-.01; -0.9[b]	-3.8	2.92	-4.4	3.23	-.02; -1.5[b]
Sp Present	-18.8	1.89	-9.7	0.88	-.04; -2.4[c]	-14.0	1.30	-17.4	1.59	-.09; -5.5[c]
Head	32.9	5.06	34.6	4.32	.18; 11.7[c]	9.0	1.24	7.8	0.95	.05; 2.9[c]
# Fam Jobs	20.3	4.44	35.2	7.15	.19; 11.9[d]	39.4	8.31	31.1	5.91	.18; 14.0[d]
Log Likel	-4231		-168627			-4689		-141054		
N	1345		1345			1668		1668		
Pve	715		715			771		771		

Notes: (*a*) Predicted probability of labour force participation F and expected hours E(H), in F; E(H) format, with continuous variables at their means and discrete variables at zero. (*b*) Effects on F; E(H) of increase in continuous variable from its mean to its maximum value. (*c*) Effects on F; E(H) of increase in dummy from 0 to 1. (*d*) Effects on F; E(H) of increase in categorical variable from 0 to 1 child. Columns 2, 4, 7 and 9, which report values for coefficients/standard errors, are in absolute terms. All equations were estimated with the same specification as in Tables 3 and 4 with the additional regressors, shown above, to control for family circumstances.