Intergenerational Income Mobility In Canada*

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Will the children of high income families achieve equally high incomes? Will the children of low income families stay in poverty? These important questions deal with intergenerational income mobility, that is with the transmission of economic status between generations. Our objective in this chapter is to determine the degree of intergenerational income mobility in Canada during the mid-1980s and 1990s and to investigate whether it has changed over time. In an era of increasing income inequality within a generation, it is important to understand whether equality of opportunity is preserved, or whether increasing polarization in labour market outcomes will be further exacerbated in the next generation. Equality of opportunity makes it possible for citizens to exploit their personal abilities and resources without consideration of family background. The higher the transmission of economic status between two generations, the lower the equality of opportunity. The degree of economic transmission between two generations, in turn, affects income inequality within a generation. For example, Becker and Tomes (1986) have shown that income mobility is affected by parental investments in their offspring, which then affects income inequality. The degree of transmission between

two generations tells us how a society deals with the issues of income inequality and equality of opportunity, it can be seen as a socioeconomic barometer of a society. It is also of great importance to many public policy concerns, such as the public financing of schooling.

Not surprisingly, many recent studies of intergenerational equity have concentrated on the two countries where within-generation income inequality has increased most dramatically: the United States (Altonji and Dunn 1991, Solon 1992, Zimmerman 1992, Peters 1992) and the United Kingdom (Deaden, Machin and Reed 1995). Another reason for the focus on these countries has to do with the availability of longitudinal data tracking individuals for sufficiently long periods. Data limitations have prevented similar research in Canada. However, Corak and Heisz (1995, 1998) have recently overcome this problem by using income tax information on close to 450,000 father-son pairs. In this chapter we use a different route to examine the same issue by combining publicly available data on average occupational income from the decennial Canadian Censuses of 1951 through 1991 (including the 1986 census) with data from the General Social Survey of 1986 and 1994. In spite of the very different data used our results are similar to those of Corak and Heisz (1995, 1998), and indicate that there is a greater degree of intergenerational income mobility in Canada than in the U.S. or the U.K. We also find a higher degree of transmission of economic status between fathers and daughters than between fathers and sons. Finally our estimates show that the rate of intergenerational income mobility is higher among more recent cohorts than among older cohorts.

1. Data Description

The choice of an appropriate data set is an especially crucial methodological issue in studies dealing with intergenerational income mobility. The need for information on the status of individuals in two successive generations at approximately the same points in the life cycle is a particularly stringent requirement. The General Social Surveys (GSS) conducted by Statistics Canada do, however, provide a rich set of information of this kind and have permitted a host of sociological studies dealing with occupational or educational mobility between generations. In particular, the respondents to the 1986 and 1994 GSSs were asked detailed questions about their parents. Respondents provided information on the employment status, the education level, the occupation and the industry of their fathers and mothers when they (the respondents) were 15 years of age. In addition, the GSS provides respondent information on gender, age, years of education, occupation and before-tax income from wages, salaries and self-employment. However, the absence of information on parental income is an important limitation for our purposes. We therefore develop a strategy to obtain a measure of incomes for the fathers using the information available on them. The resulting estimation strategy is equivalent to an instrumentation of fathers' incomes with fathers' occupations and implies problems of its own, which we discuss in detail in Section 2.

Another critical issue in the estimation of intergenerational mobility is the determination of parental permanent income. Solon (1992) and Zimmerman (1992) have shown that the use of income in a single year can seriously overstate the degree of income intergenerational income mobility, especially in combination with the use of an overly homogeneous sample. We use the occupational information from the GSS to obtain

average income by occupation from the Census tables under the assumption that occupation is a good instrument to estimate the parent's permanent income. In the 1986 GSS the occupation of the parents is given only in terms of the Pineo-Porter-McRoberts Socioeconomic Classification of Occupations (Pineo, Porter and McRoberts 1977). This system reclassifies the Canadian Standard Classification of Occupations (4-digit codes) of 1971 and 1981 into 16 categories, ordered by skill levels. These range from farm labourers, unskilled manuals, and unskilled clerical sales and services; to semi-skilled manuals, semi-skilled clerical sales and services, farmers; to skilled crafts and trades, skilled clerical sales and services, foremen, supervisors, middle managers, technicians, semi-professionals; and finally to high level managers and professionals. The ordering by skill levels represents the main advantage of the Pineo-Porter-McRoberts reclassification over the standard 2-digit reclassification of the 1980 occupation codes provided by Statistics Canada. The latter was not designed to measure economic status and sometimes aggregates occupations by industry rather than by skill levels. In fact the new 1990 occupation codes have been completely redesigned to correct this problem. The Pineo-Porter-McRoberts system is relatively well suited to our needs since it can be applied to the earlier classifications systems (Pineo, 1985). Nonetheless, this reclassification system presents some problems. The disadvantage of an ordering by skill levels is that the resulting classes may represent a job ladder rather than lifetime occupational choices. Clearly, an improvement in our estimation strategy would be to use the detailed occupation codes and obtain average occupational income by years of education. However, such a strategy is not feasible from publicly available data.

Since detailed 4-digit occupation codes are not available in the Census public release files, we use the average employment income by detailed occupation from the "Tables of Average Employment Income by Detailed Occupation." These tables offer the average income of all workers 15 years old and older, by gender for the Census years 1951 to 1991. For every Census table, we re-classify every 4-digit occupation by the Pineo-Porter-McRoberts classification and compute the weighted average employment income for the 16 occupational categories. The estimated average occupational income is then the weighted average occupational income for 1950, 1960, 1970, 1980, 1985 and 1990. Fathers' occupations are observed when the respondents were 15 years of age: from 1944 to 1982 for the 1986 sample, and from 1952 to 1990 for the 1994 sample. We linearly interpolate between every two Census periods to obtain the intermediate incomes. The estimated average occupational income computed in constant 1993 dollars are shown in Table 4.1. The mean income of the 16 occupational groups increased from $15,434 in 1950 to $34,979 in 1980, then decreased to $33,245 in 1985, and went back up to $34,251 in 1990. Alongside this rising trend in mean income is an increase in income inequality, the gap between the highest and lowest income widening substantially over the period. In 1950 the average income of unskilled manual workers was about half of that of high level managers, while in 1990 it was less than a third.

Our measure of occupational income averages the income of workers at different points in their life-cycle, and therefore provides a better measure of economic status than one observation at a point in time. Income observed at any one year will contain both a permanent and a transitory element. Solon (1992) and Zimmerman (1992) average up to five consecutive years of father's income data in order to minimize the problems of biases

arising from the presence of transitory components. (An ideal measure of permanent income would be based upon a series of observations over an individual's entire life.) Our average occupational income measures may be contaminated by transitory components only insofar as these measures are linked to a transitory occupation. The issue of whether the father's occupation, as reported by the sons, is likely to have been a permanent or a transitory occupation, may depend on the age of the father. The older the father the more likely the occupation is permanent. A related issue is whether our measures of income are contaminated by life-cycle effects. Is the average income of high-level managers higher because of economic status, because high-level managers are further along their life-cycle paths, or because their life-cycle path is steeper? We cannot assess age differences by occupation for fathers, but the average age of sons from the 1986 and 1994 GSS does vary somewhat by the Pineo-Porter-McRoberts occupations. High-level managers and foremen are on average four years older than middle managers and skilled craft workers. The average age differences between the other occupations are smaller. One way to address the issue of differences in life-cycle paths by occupation would be to use different life-cycle corrections for each occupation. While this strategy is feasible for children, it is not for fathers and thus simply reduces the correlation between father and child incomes. An alternative strategy is to compare fathers and children at the same point in their life-cycle paths, assuming that these paths have not changed significantly over time. We do not know the age of the father, but if we assume 25 to 35 year differences between fathers and sons, this places the fathers of 40 to 50 year old children at the same stage in the life-cycle as their children. Thus estimates that focus on this age group are less likely to be contaminated by problems associated with transitory occupations and life-cycle effects.

Our samples from the GSS consist of men and women aged 17 to 59 years, whose main activity in the 12 months prior to the survey was working at a job or being self-employed. The main characteristics of the samples used in the estimation are presented in Table 4.2. We do not study intergenerational income mobility between mother-son and mother-daughter pairs mainly because we have information on the mother's occupation only for a fairly small sample. Our sample sizes are 3,400 father-son pairs and 2,474 father-daughter pairs from the 1986 GSS, and 2,459 father-son and 2,308 father-daughter pairs from 1994 GSS. The men in our sample are on average 35.5 years of age in 1986 and 37.4 in 1994, while the women average 34.2 years of age in 1986 and 36.5 in 1994. Overall the women are more educated than the men. In addition, fathers are less educated than their children. The estimated occupational weekly income for sons is lower than the reported average income in the 1986 GSS; this discrepancy is much less important in the 1994 GSS. It is difficult to know whether this reflects life-cycle or cyclical effects.[1]

Our sample size compares favourably to recent U.S. studies, which have been based on as few as 348 to a maximum of 876 observations (Altonji and Dunn 1991, Solon 1992, Zimmerman 1992). For the U.K. Atkinson (1981) uses a sample taken from a study of men in York in 1950 and traces the sons of these individuals during the late 1970s. His final non-random sample covers only 307 father-son pairs. However, another advantage of our data is the information it contains on individuals in many more cohorts. Because we have two observations on the same cohorts, we will also be able to perform some cohort analysis to investigate whether intergenerational mobility has changed over time. In

[1] Note that the Census tables average income from individuals aged 15 and up. Because employment among young people went down from 1986 to 1990, it is possible that the average age of that sample went up, thus mimicking our sample more closely.

contrast, Dearden, Machin and Reed (1995) use the National Child Development Survey, an ongoing survey of all persons born in the U.K. between March 3rd and 9th of 1958. They have access to child incomes only at age 23 and 33 and have only a single measure of father incomes when the children were 16 years old. In Canada, Corak and Heisz (1995, 1998) focus on a cohort of 27 to 31 year olds using income drawn from tax records and multi-year averages for the father's income.

In our analysis we follow the thrust of the existing literature and appeal to two complementary analytical frameworks: the log-linear regression model and the quartile transition matrix. These methods have been applied by many authors and make comparisons possible. The log-linear regression model posits a simple linear relationship between the child's and the father's log income and may be interpreted as examining the average degree of transmission of economic status. The quartile transition matrix attempts to distinguish different degrees of transmission by income quartile. It allows us to assess whether there is more or less mobility at the bottom than at the top of the income distribution.

2. Regression Models

In the log-linear regression model the logarithm of the permanent income of a child belonging to family i (y_i^{child}) is considered to be a linear function of the log of the permanent income of the father (y_i^{parent}) as shown by equation (1):

$$y_i^{child} = \alpha + \beta y_i^{parent} + \varepsilon_i \qquad (1)$$

where ε_i is an error term usually assumed to be distributed as $N(0, \sigma^2)$. This equation should be viewed as a reduced-form equation of a complex process of economic

transmission, where the coefficient β indicates the degree of mobility between the two generations. There are two extreme cases. First if β=0, there is complete mobility, in other words complete regression to the mean. The income of the child shows no correlation with the father's income. At the other extreme, if β=1, there is complete immobility. The distribution of income in the father's generation is completely preserved in the child's generation. For values of β between 0 and 1, there is regression towards the mean, but the rate depends on the value of β. The lower β, the lower the chances the child will inherit the economic status of his father, and the higher the degree of mobility.

Methodological problems arise because we do not observe the child's or the father's permanent income, and have to use proxies for it. As mentioned earlier, Solon (1992) and Zimmerman (1992) have pointed out that an error-in-variables problem may arise from the use of proxies that include transitory components. While these transitory components will generally be uncorrelated with permanent status, they will lead to an overestimation of the variance of the father's income. This results in an underestimation of the true intergenerational transmission coefficient (which is equal to the covariance between the child's and the father's income divided by the variance of the father's income). We do not face problems arising from an overstated variance of the father's income. In fact, our measure of father's income may understate the true population variance which, in itself, would lead to estimates of β that are upwardly biased. However, this smaller dispersion of father's income may also lead to smaller covariance between the child-parent incomes. Therefore, it is difficult to assess the direction of the overall bias in our models.

We use the detrended average occupational employment income of fathers during the years the child was 15 years old. This measure corresponds to the instrumentation of the father's income with the father's occupation, and may be a subject of concern. As explained in Solon (1992), if the father's occupation is part of a structural model of the son's income, the corresponding instrumental variable estimate of β may be biased. If the father's occupation does not influence the son's income beyond its indirect effect through the father's income the estimate will be consistent. However, it will be upwards or downwards inconsistent if the father's occupation influences (positively or negatively) the son's status beyond its effect through income levels. We will thus perform an overidentification test to verify that our estimates are consistent. For children, we have information on employment income in 1985 from the 1986 GSS and in 1993 from the 1994 GSS. This single measure of income may deviate from permanent status because of age effects that we attempt to eliminate by using time-varying control factors that may affect current income (specifically age and age^2). We will also provide alternative estimates using the same instrumented measure of income for sons as we do for fathers. While we are aware of the potential problems with our measure of permanent income, we believe that these problems may not have evolved differently across age groups or over time and thus that our age groups and cohort analyses will be informative.

Two different approaches are used to estimate β. The first—which we refer to as the Instrumental Variables (IV) approach—assumes that occupation is a valid instrument to estimate the permanent income of the father. If this is the case then the father's income can be expressed as:

$$y_{it}^{parent} = \quad_{kt} \quad Occ_{kt}^{parent} + v_{it}^{parent}, \qquad (2)$$

where y_{it}^{parent} is the occupational income of the father in year t, Occ_{kt} are a set of dummy variables indicating occupation k (k=2 to 16) in year t, $_{kt}$ is the average income of occupation k in year t, and v_{it} is an error term. We recover a residual measure of the permanent status of the parent by removing the time varying components from the occupational income,

$$_{ri}^{parent} = _{kt} \; Occ_{kt}^{parent} - \; x_t^{parent}, \qquad (3)$$

where $_{kt}$ is the estimated average income for occupation **k** in year **t** ,where the time varying factors $\mathbf{x}_t^{parent}$ are dummies linked to the year of observation of the father's occupation.

Similarly, we do not observe the child's permanent income y_i^{child} , but rather y_{it}^{child}, that is employment income at a single point in time. We assume that employment income at time t is a function of the permanent income y_i^{child} and also of x_{it}^{child}, a vector of observed time-varying factors which affect the current status of the adult "child" and of w_{it}^{child}, a transitory error term:

$$y_{it}^{child} = y_i^{child} + \; x_{\;it}^{child} + \; w_{it}^{child} \qquad (4)$$

To obtain a measure of permanent income, we regress y_{it}^{child} on the time-varying factors x_{it}^{child} (age, age^2) and use the residual as an estimate of the child's permanent income:

$$_{ri}^{child} = \; y_{it}^{child} - \; x_{it}^{child} = \; y_i^{child} + w_{it}^{child} \qquad (5)$$

When we run the following regression, we have thus removed the time-varying factors that we can control for,

$$_{ri}^{child} = \; + \; _{ri}^{parent} + \; _{i}, \qquad (6)$$

so that

$$\mathbf{y_i^{child}} = \; + \; (\; _{kt} \; \mathbf{Occ_{kt}^{parent}} - \; \mathbf{x_t^{parent}}) + \; _{i}^{child}, \qquad (7)$$

where $_i^{child} = \quad _i + \mathbf{W}_{it}^{child}$. The OLS estimator of the coefficient will be consistent if the occupation dummies, $\mathbf{Occ}_{kt}^{parent}$, and the trend variables, $\mathbf{x}_t^{parent}$, are uncorrelated with $_i^{child}$.

The second approach—referred to as the Occupational Income (OccInc) approach—assumes that the transmission processes between permanent income and occupational income is the same, that is, the essential features of economic status are transmitted through occupational status. This implies that the correlation between father and child permanent incomes will be the same as between father and child occupational income:

$$\text{corr}(y_i^{child} , y_i^{parent}) = \text{corr}(_{oi}^{child} , \quad _{oi}^{parent}). \qquad (8)$$

If our assumption is correct, we can estimate the following regression ,

$$_{oi}^{child} = \quad + \quad _{oi}^{parent} + e_i, \qquad (9)$$

where $_{oi}^{child} = \quad _{kt} \quad Occ_{kt}^{child} - \quad x_t^{child}$ and $_{oi}^{parent} = \quad _{kt} \quad Occ_{kt}^{parent} - \quad x_t^{parent}$, and where $_{kt}$ is the estimated income of occupation k in year t and where the time-varying controls are the same as in (3). If the assumption about the similarity of the transmission of occupational income and permanent income is valid, and if the occupational dummies and trend variables are uncorrelated with e_i, the coefficient estimated through equation (9) will be a consistent estimate of the true coefficient of transmission of economic status.

Table 4.3 offers the results from the log-linear model for fathers and sons and fathers and daughters, separately for the two surveys and for the joint sample. We also conduct the analysis separately for three different age groups. This allows us to compare our estimates to other studies that focus on particular age groups. The estimates of the coefficient of intergenerational mobility (what we refer to as) using the IV approach are given in Panel A. For the 1986 GSS we estimate to be 0.191 for father-son pairs,

and 0.228 for father-daughter pairs. These estimates are essentially the same when the 1994 data are used: 0.217 and 0.226.[2] Overidentification tests are performed to find out if there remains some correlation between the error term $_i$ and the dummy variables Occ_{kt}. For the 1994 sample of sons and daughters, the regression of the residuals from equation (7) on the variables Occ_{kt} passes the F test of non-significance. We conclude that our instrumentation strategy seems valid for the 1994 sample. On the other hand, the 1986 sample for sons and daughters almost passes the F test at 10%, but the null hypothesis is rejected; the two terms are correlated but the correlation is not strong. The fact that the estimates from 1986 and 1994 are not significantly different further indicates that the biases are quite small.

The results from the regressions that split the data into three age groups are presented in Panel B of Table 4.3. Intergenerational mobility diminishes with age for sons in the 1986 and 1994 samples. The estimates of for the father-son pairs from the 1986 GSS are 0.105 when the sons are between 17 and 29 years old, 0.201 when they are between 30 and 39 years, and finally 0.297 when they are 40 to 59 years old. The same pattern is found for the father-daughter pairs of the 1994 sample, but with 1986 data the group showing the least mobility are 30 to 39 years of age. Interestingly, this group represents approximately the same cohort of women aged 40 to 59 in 1994. This raises the possibility of cohort effects. The higher mobility of the older group in 1986 may reflect a change in the work patterns of women, or may simply be attributed to smaller sample sizes. The sons' and daughters' income are adjusted for life-cycle effects by using the

[2] The differences between the 1986 and 1994 results are not statistically significant The reported standard errors, however, do not take into account the fact that the measure of the father's income is an estimate and thus probably overstate the true standard errors.

residuals of the regression of their income on their age and age square. It is possible that this adjustment does not fully capture the life-cycle effects. Alternatively, intergenerational mobility could have been different for different cohorts. The other interesting implication of these results for studies concentrating on children in their early 30s is that this choice for father-son pairs is representative of the whole sample.

The results from the occupational income approach are offered in Panel C of Table 4.3. To repeat, this approach assumes that the transmission process between generations is the same for permanent incomes as for occupational incomes. For the sons, the estimates of for the 1986 sample and 1994 sample are respectively 0.185 and 0.202. These estimates are not significantly different from those based on the IV approach. However, they are significantly lower for daughters: 0.155 for 1986 and 0.139 for 1994 (versus 0.228 and 0.226).[3] The occupational income results are in line with other studies that have found somewhat smaller coefficients with a prediction approach (Dearden, Machin and Reed 1995). They also indicate that our estimates for father-son pairs are more robust than our result for daughter-father pairs.

Another way to split our samples is by birth cohort (the approximate year of birth). An analysis of this sort reveals whether the degree of intergenerational income mobility has changed over time. Table 4.4 presents results for three cohorts. Individuals aged 50 to 59 in 1986 were removed from the joint sample in order to have two observations over time for each cohort. The first cohort consists of individuals born between 1935 and 1945 (who were in their forties in 1986 and in their fifties in 1994); the second consists of

[3] The estimated average occupational income for the daughters sample are average occupational income from a women sample.

individuals born between 1946 and 1954 (those in their thirties in 1986 and in their forties in 1994); and the third of individuals born between 1955 and 1969 (in their twenties in 1986 and their thirties in 1994). The findings in Table 4.4 are based on the IV approach.

There are significant differences between the three cohorts. For the father-son pairs, we find an estimate of 0.316 for the first cohort, 0.246 for the second cohort and 0.157 for the third. Because we have only two observations over time, there remains substantial age differences between the cohorts, and for this reason our results showing an increasing mobility over time for sons should be interpreted with caution. We observe a different pattern for the father-daughter pairs. The estimates are 0.265 for the first cohort, 0.323 for the second and 0.191 for the third. The second cohort is the one showing the highest degree of transmission of economic status. The differences in participation rates of women across these cohorts may be at the origin of these different results for the father-daughter pairs. This second cohort represents the first generation of women to enter the labor market in great numbers. The degree of transmission of economic status between two generations seem to differ across cohorts, but it is not clear whether this is due to age differences, differences in the transmission process (for example, through universal access to higher education), or to an increase in the dispersion of income for younger cohorts. For example, since the coefficient of transmission of economic status is equal to the coefficient of correlation between two generations times the ratio of the standard deviations of fathers and children's incomes, a relative increase of 20% in the standard deviation of children's income would lead to an increase of 0.03 in if the coefficient of correlation was 0.3.

In summary, our main estimates of the degree of intergenerational income mobility are in the range of 0.2, the consensus value that was found in the earlier studies using simple least squares regression models. These earlier estimates were criticized as being downwardly biased because the observed father's income was thought to include transitory components. While the presence of transitory components is less likely to be a problem with our estimation strategy, the fact that the variance of the father's occupational income may understate the true variance of the father's permanent income may be problematic. It is thus interesting to compare our results with findings of other recent studies conducted in the U.S., the U.K., and especially Canada. Some results from this literature are presented in Table 4.5. We find that our estimates of the degree of intergenerational income mobility (ranging from 0.19 to 0.21) between sons and fathers are very similar to those of the other Canadian study. Corak and Heisz (1995) obtain an estimate of 0.191 for the father-son pairs using a five years average of fathers' income with Least Squares regression. Their other estimates are even lower. A more conservative interpretation of our results would focus on the forty to fifty year group, and place our estimate of income mobility in the low 0.3 range. Conversely, the latest studies done in the U.S., by Solon (1992) and Zimmerman (1992), show estimates of between 0.413 and 0.538 using five and four years average of fathers' incomes with Least Squares regressions. Similarly, the estimates for U.K. males found by Atkinson (1981) and Dearden, Machin and Reed (1995) are substantially higher than those found for Canada by Corak and Heisz (1995, 1998), and by us for similar age groups. Of course, there remain important differences in methodology. However, even when we compare the admittedly flawed Least Squares estimates presented in Solon (1992), Zimmerman (1992), and

Dearden, Machin and Reed (1995) (and reported in Table 4.5) to the Canadian estimates (including our findings) these are still lower than those found in the U.S. or in the U.K. We thus conclude that there is more intergenerational income mobility in Canada than the U.S. and the U.K. It is interesting to note that Björklund and Jäntti (1997) in a meta-study reach a similar conclusion: the U.S. and the U.K. have the lowest mobility among seven other developed countries. We will now attempt to corroborate this conclusion using another technique to estimate the degree of transmission of economic status between generations.

3. The Transition Matrix Method

The quantile transition matrix method allows us to analyze movement between income quantiles across generations. Since it is a widely used method it is possible to make comparisons with other studies. The major advantage of quantile transition matrices over the log-linear regression model is that it permits an assessment of whether there is more or less mobility at the bottom or at the top of the income distribution. If one is preoccupied by the existence of an "underclass," the degree of mobility in the lowest quantiles of the income distribution are more relevant than the average degree of intergenerational mobility. One should note, however, that because our study is limited to earnings mobility, it does not address issues related to the existence of a "poverty trap." There are generally no earners in Canadian families ranked in the lowest decile of family income to need ratios (Fortin and Lemieux, 1997). In addition one must keep in mind that quantile transition matrices provide only a very crude way of looking at potential non-linearities in the

transmission process. For example, Corak and Heisz (1998) examine this issue in much more detail using non-parametric regression methods.

The methodology of transition matrices is the following. The fathers and the children are each ranked according to income, and divided (in our case) into four groups of equal size. Individuals in the first group have the lowest income and those in the fourth group the highest. A matrix is then constructed where each element, referred to as a_{ij}, represents the probability that a child will be in quartile j if his or her father was in quartile i. This matrix exhibits a bi-stochastic property. If a_{ij} is the proportion of children of fathers in quartile i who enter quartile j, then $\sum_i a_{ij} = 1$ and also $\sum_j a_{ij} = 1$. The two extreme cases of income mobility can be represented using this approach: complete mobility occurs when each element of the matrix equals 0.25; complete immobility when the diagonal elements are equal to 1 (and all the others equal 0).

The estimates of the transition matrices for 1986 and 1994 (for sons and daughters) are given in Table 4.6. These are computed using the age-corrected residual income for the sons and daughters and the detrended average occupational income for fathers. Generally, the results display a high rate of mobility. Even in the top quartile, where there is usually less mobility, the rates of transmission of economic status are lower than those found in other studies. For the 1986 and 1994 samples, the probabilities that sons whose father were in the top quartile of the income distribution stayed at the top are 0.32 and 0.33. These correspond to the estimates in Corak and Heisz (1995), but are much lower than the 0.40 for the U.S. reported by Peters (1992) or the 0.39 for the U.K. reported by Dearden, Machin and Reed (1995). The same findings apply to the daughter samples, with estimates of 0.34 for 1986 and 0.33 for 1994. The other notable finding, by comparison

with other studies, is that the mobility at the bottom of the income distribution is quite substantial. In fact, the 1994 results tell us that sons whose fathers were in the bottom quartile of the income distribution are more likely to be in the second or third quartile than in the bottom one. Our estimated probability that sons whose fathers were in the bottom quartile remain in that quartile is 0.260 to 0.285, somewhat lower than the Peters and Dearden et al. findings for the U.S. (0.42) or the U.K. (0.315). While our estimates are generally not statistically different from 0.25, Corak and Heisz (1995) found a larger estimate (0.353) that is statistically different from the situation of perfect mobility. Corak and Heisz (1998), however, show that finer disaggregation (the use of deciles or percentiles) pushes any stickiness in the transmission process toward the very low and very high income classes. The general pattern of income mobility is thus relatively close to the situation of perfect mobility, with the exception of moves from either the very top or the very bottom of the income distribution.

The ranking of different quartile matrices can be useful to compare the degree of mobility found in the various studies. An Immobility Index can be computed as the ratio of the sum of the diagonal elements to the sum of all elements in the matrix. We find an immobility index of 0.28 for the father-son pairs and 0.27 for the father-daughters pairs, as shown in Table 4.6. These estimates are in the range of those of Corak and Heisz (1995), who obtained an index value of 0.306 for Canada. As the authors report, these values are substantially lower than those found by others for the United Kingdom (0.3675) and the U.S. (0.350). These findings support the results of our log-linear model, and suggest that if intergenerational income mobility is higher in Canada than in the U.S. or in the U.K.,

this would be mainly because of more mobility both at the bottom and at the top of the income distribution.

4. Conclusion

In this chapter we combine publicly available data on average occupational income from the Canadian Censuses of 1951 to 1991 with data from the General Social Surveys of 1986 and 1994 to obtain estimates of intergenerational income mobility. We use an instrumental variables approach to the estimation of the standard log-linear models, as well as some cohort analyses of these data. Our estimates of the log-linear model are similar to the ones found by Corak and Heisz (1995, 1998), the only other Canadian studies done to date. These results suggest that intergenerational mobility is higher in Canada than in the U.S. or the U.K., where recent research has shown that previous estimates overstated the degree of mobility. Because of our somewhat larger sample sizes, we are also able to perform an analysis by birth cohort. This shows an increase in the degree of intergenerational income mobility over time. Income mobility is also greater for younger age groups. Finally, we construct quartile transition matrices that suggest intergenerational income mobility is higher in Canada than in the U.S. or the U.K. because there is more mobility at the top and (to a lesser extent) at the bottom of the earnings distribution. To drawn implications for public policy, the next step would be to investigate the dynamic process behind the transmission of economic status between two generations.

There are some important caveats that must be kept in mind when assessing our results. Since detailed occupation codes are not released in the public use files of the General Social Surveys, we had to estimate average occupational income for only 16

occupational groups. It is difficult to know whether this averaging process over detailed occupation categories and individuals of different ages brings us closer to an ideal estimate of the father's permanent income than estimates that average individual income for a few years. The best test of our methodology would be to apply it to the same U.S. data used in other studies and compare the findings.

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Table 4.1
Estimated Average Occupational Income of Men in 1993 Dollars

Pineo-Porter-McRoberts Classification of Occupations	1950	1960	1970	1980	1985	1990
			(1993 Dollars)			
Professionals	20,888	32,457	43,547	54,153	53,928	54,872
High level management	23,223	36,734	81,227	71,378	69,702	68,871
Semi professionals	18,293	27,100	31,518	35,753	34,628	36,632
Technicians	16,122	24,178	29,401	35,190	35,361	36,207
Middle management	15,793	18,194	44,405	46,226	42,607	45,425
Supervisors	17,009	25,696	34,208	36,398	34,894	36,192
Foremen	19,851	26,033	37,596	42,217	39,335	40,763
Skilled clerical sales and service	16,713	23,087	31,851	37,661	35,856	36,380
Skilled crafts and trades	15,381	20,852	27,941	32,205	30,835	32,155
Farmers	11,714	17,258	15,568	25,589	19,772	20,190
Semi skilled clerical sales and service	14,653	16,581	21,048	22,478	21,209	22,052
Semi skilled manuals	13,039	18,326	23,859	27,058	25,886	26,445
Unskilled clerical sales and service	11,701	17,671	20,201	23,483	21,736	21,976
Unskilled manuals	11,549	14,197	18,736	22,143	21,093	22,030
Farm labourers	5,576	7,437	9,637	12,744	11,825	13,569

Source: Calculations by authors from Statistics Canada data.

Table 4.2
Descriptive Statistics

	Sons		Daughters	
	1986	1994	1986	1994
Age	35.5	37.4	34.2	36.5
	(10.12)	(10.76)	(9.98)	(10.70)
Years of education	12.9	13.85	13.32	14.09
	(3.08)	(2.98)	(2.62)	(2.57)
Weekly Employment Income	$776.83	$702.94	$521.83	$474.48
	(456.86)	(383.50)	(365.79)	(284.59)
Occupational Weekly Income	$641.35	$680.80	$362.23	$398.58
	(241.83)	(245.50)	(143.30)	(151.42)
Father's Education	9.2	10.5	9.3	10.6
	(4.66)	(5.07)	(4.44)	(4.97)
Father's Occupational Weekly Income	$493.08	$559.83	$512.06	$560.71
	(223.61)	(250.45)	(228.51)	(237.26)
Sample size	3,400	2,459	2,474	2,308

Notes : All incomes are in constant 1993 dollars. The numbers in parentheses are standard deviations.
Source: Calculations by Authors from General Social Survey, Statistics Canada.

Table 4.3
Estimates of the Coefficient of Intergenerational Mobility

	Fathers and Sons		Fathers and Daughters	
	1986	1994	1986	1994
A. Instrumental Variables Method				
	0.191	0.217	0.228	0.226
	(0.029)	(0.032)	(0.041)	(0.040)
	[3400]	[2459]	[2474]	[2308]
B. Instrumental Variables Method by Selected Age Groups				
	0.105	0.048	0.143	0.145
17 to 29 years	(0.052)	(0.065)	(0.062)	(0.070)
	[1103]	[651]	[959]	[681]
	0.201	0.218	0.324	0.218
30 to 39 years	(0.043)	(0.051)	(0.063)	(0.064)
	[1220]	[811]	[842]	[733]
	0.297	0.351	0.208	0.309
40 to 59 years	(0.062)	(0.055)	(0.104)	(0.076)
	[1077]	[997]	[673]	[894]
C. Occupational Income Method				
	0.185	0.202	0.155	0.139
	(0.017)	(0.021)	(0.020)	(0.024)
	[4013]	[2335]	[3027]	[2153]
D. Joint Sample, Instrumental Variables Method				
	0.208		0.228	
	(0.022)		(0.029)	
	[5859]		[4782]	

Note: ( ) indicates standard errors, [ ] indicates sample size.

Table 4.4
Estimates of the Coefficient of Intergenerational Income Mobility Using Cohort Analysis

	Fathers and Sons	Fathers and Daughters
Cohort born between 1935 and 1945	0.316 (0.065) [1061]	0.265 (0.099) [756]
Cohort born between 1946 and 1954	0.246 (0.035) [1827]	0.323 (0.050) [1411]
Cohort born between 1955 and 1969	0.157 (0.034) [2248]	0.191 (0.041) [2021]

Note: ( ) indicates standard errors, [ ] indicates sample size.

Table 4.5
Estimates of Intergenerational Income Mobility from other Studies

Author	Country and Data Set		Estimate of for father-son pairs
Corak and Heisz (1995)	Canada, Tax Record Data (1992) 450,000 father-son pairs sons aged 28 to 31	OLS with single year of father's income	TT: 0.121-0.136 AE: 0.115-0.143
		OLS with five years average of father's income	TT: 0.191 AE: 0.172
Altonji and Dunn (1991)	United States, NLSY (1965-1967) 678-739 father-son pairs sons aged 29 to 39	OLS with time averaging of father's income and age controls	AE: 0.180 HW: 0.263
		IV with later years income as instruments and full set of controls	AE: 0.218 HW: 0.282
Solon (1992)	United States, PSID (1984) 348 father-son pairs sons aged 25 to 33	OLS with single year of father's income	AE:0.386 HW:0.294
		OLS with five years average of father's income	AE: 0.413
		IV with father's years of education	AE: 0.526 HW:0.449
Zimmerman (1992)	United States, NLSY(1981) 876 father-son pairs sons aged 29 to 39	OLS with four years average of father's income	AE:0.538 HW:0.391
		IV with Duncan index for father's status	AE: 0.417 HW:0.485
		IV with forward quasi-difference instrument	AE: 0.36 HW:0.379

Atkinson (1981)	United Kingdom, Rowntree Survey (1975-1978) 288-307 father-son pairs sons aged 25 up	OLS	WE:0.358 HW:0.428
		OLS with life-cycle adjustments	HW:0.415
Dearden, Machin and Reed (1995)	United Kingdom, NCDS(1991) 1665 father-son pairs, 747 father-daughter pairs children aged 23 and 33	OLS with single year of father's income	WE:0.216 (sons) WE:0.352 (daughters)
		IV with father's education and social class	WE:0.581(sons) WE:0.669(daughters)
		Predicted wages with father's education and social class	WE:0.425(sons) WE:0.469(daughters)

Note: PSID - Panel Study of Income Dynamics; NLSY - National Survey Study of Youth; NDCS - National Child Development Study; TT - Total Income; AE - Annual Earnings; WE - Weekly Earnings; HW - Hourly Wage.

Table 4.6
Quartile Transition Matrices

A. Fathers – Sons, 1986

		Sons			
		Bottom	2nd	3rd	Top
Fathers	Bottom	0.285 (0.016)	0.279 (0.015)	0.234 (0.015)	0.224 (0.014)
	2nd	0.293 (0.016)	0.239 (0.015)	0.212 (0.014)	0.216 (0.014)
	3rd	0.232 (0.014)	0.258 (0.015)	0.256 (0.015)	0.252 (0.015)
	Top	0.184 (0.013)	0.222 (0.014)	0.296 (0.016)	0.322 (0.016)

Immobility Index=0.276

B. Father – Sons, 1994

		Sons			
		Bottom	2nd	3rd	Top
Fathers	Bottom	0.26 *(0.018)*	0.277 *(0.018)*	0.264 *(0.018)*	0.19 *(0.016)*
	2nd	0.265 *(0.018)*	0.251 *(0.018)*	0.234 *(0.017)*	0.221 *(0.017)*
	3rd	0.254 *(0.017)*	0.213 *(0.016)*	0.278 *(0.017)*	0.303 *(0.018)*
	Top	0.187 *(0.015)*	0.241 *(0.017)*	0.231 *(0.018)*	0.332 *(0.019)*

Immobility Index = 0.280

C. Fathers – Daughters, 1986

		Daughters			
		Bottom	2nd	3rd	Top
Fathers	Bottom	0.265 (0.018)	0.272 (0.018)	0.278 (0.018)	0.196 (0.016)
	2nd	0.276 (0.018)	0.248 (0.018)	0.231 (0.017)	0.228 (0.017)
	3rd	0.28 (0.018)	0.23 (0.017)	0.232 (0.017)	0.257 (0.017)

Top	0.171 (0.015)	0.233 (0.017)	0.276 (0.017)	0.338 (0.019)

Immobility Index = 0.271

D. Father –Daughters, 1994

		Daughters			
		Bottom	2nd	3rd	Top
Fathers	Bottom	0.256 (0.018)	0.274 (0.019)	0.244 (0.018)	0.22 (0.017)
	2nd	0.281 (0.019)	0.27 (0.018)	0.239 (0.018)	0.211 (0.017)
	3rd	0.25 (0.018)	0.229 (0.018)	0.246 (0.018)	0.281 (0.019)
	Top	0.187 (0.016)	0.236 (0.017)	0.25 (0.018)	0.326 (0.019)

Immobility Index = 0.275

Note: Child incomes are the age-corrected residual incomes, and father incomes are the detrended average occupational incomes.
( ) represents the standard errors due to sampling variability and should be interpreted as a lower bound to the true standard errors.