This paper uses longitudinal tax data to study the impact of growing up in a lone parent and/or low income family on a young woman’s probability of herself being a lone parent and/or being in low income status one generation later. Neighbourhood characteristics – the percentage of lone parent and low income families – are also considered, allowing us to decompose the child’s environment into a family effect and a neighbourhood effect. Our preferred econometric specification is a bivariate probit model, which allows for inter-independence of the two outcomes. We find that family effects generally matter the most to both outcomes, but neighbourhood effects, particularly the lone mother variable, also matter, especially for the lone-mother outcome. Both sets of background effects matter most in the late teen years, but earlier childhood experiences also have significant (independent) effects, with some nonlinearities found in this regard (i.e., the pre- and early-adolescence years matter least of all).
I. INTRODUCTION

Census data show that lone parent families accounted, in 2001, for about 15 percent of all Canadian families, the vast majority of these (81%) headed by women. These families are among the most vulnerable in society, with about 40 percent of lone mother families living in a state of low income in 2001, as opposed to 25 percent for all families taken together, while their low income spells also tend to be longer and deeper than others. Because of the hardships they suffer, their draw on public resources, the potential handicaps their children face, and other related issues, lone mother families, especially those in low income, continue to draw considerable attention from policy makers and social scientists alike.

The contribution of this paper is to add to our understanding of lone mothers and their associated low income status by reporting the results of an empirical analysis of the impact of growing up in a lone parent and/or low income environment on a young woman’s probability of herself being a lone parent and/or in low income status one generation later. Any improved understanding of the process of entering lone-mother-low-income status is interesting not just for academic reasons, but also because it could contribute to the development of preventative or remedial measures, thus potentially reducing the magnitude of the associated problems.

For this analysis, we employ the unique properties of Statistics Canada’s Longitudinal Administrative Databank (LAD), a 20 percent sample of Canadian tax filers. In the first stage, we link young women to their families of origin in their teen years (when they start to file income tax forms), and then estimate bivariate probit models of the joint probabilities that the young woman goes on to be a lone mother or to be in low income. The two key sets of explanatory variables in these models are i) whether the woman came from a lone mother family or one that was in low income and ii) whether the family lived in a neighbourhood characterised by greater percentages of lone parent or low income families. We thus estimate the intergenerational transmission of lone mother and low income status while separating the woman’s earlier environment into family and neighbourhood effects.

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1 In this paper the term “in low income” is used to indicate that a person has an income below a specified low income cut-off – analogous to being “in poverty”. (Canada has no official poverty line and the “low income” term is typically used instead.)

2 See Finnie and Sweetman (2003) regarding the poverty dynamics of Canadian lone mother families and related policy discussions, and Dooley (1999) for a profile of lone mothers’ participation in welfare programs.
In the second stage we use the established daughter-mother links to connect to the mothers’ own tax records in the approximately 20 percent of (random) cases these also appear on the LAD to follow the daughter’s situation back in time through the parent. We then estimate the lone-mother-low-income models including the effects of the family and neighbourhood influences at different points over the daughter’s childhood.

Overall, we find that a woman’s earlier family characteristics are strong and significant determinants of her own low income and lone mother status and that the neighbourhood characteristics also play a role – but different factors affect the low income and lone mother outcomes in different ways. The data also show that both sets of background effects matter most in the late teen years, but that earlier childhood experiences also have significant effects, while the earlier and later stages seem more critical than the pre- and early-adolescent periods, thus reflecting a “U” pattern in terms of when background matters most.

Given that the existing literature – mostly U.S. – shows mixed findings for a roughly comparable set of inter-generational transmission processes between individuals’ family backgrounds and welfare participation and lone motherhood, the uniformity and strength of the family effects found here is significant. Furthermore, the neighbourhood effects add a new dimension to these transmission processes and the identification of the clear variation in effects over the different periods of a child’s development is notable. Various implications of the findings are discussed.

II. THE LITERATURE

The existing literature identifies three main factors that might underlie the inter-generational transmission of lone mother or low income status. The first is investments in human capital: low income or lone mother families are likely to have fewer resources to invest in their children’s development, while the neighbourhood environment, including local school quality, might not provide the same opportunities as those available to children from higher income or two-parent families.3 Secondly, there may be an inter-generational transmission of preferences, or what are sometimes in this literature called “imitation effects”. Girls raised in lone mother families or in

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3 See Becker and Tomes (1979, 1986) for these ideas placed in the context of a parental consumption-investment model.
neighbourhoods with a relatively high percentage of lone mothers may, for example, perceive less “stigma” from becoming lone mothers themselves, and those in low income families or in low income neighbourhoods might be influenced in comparable ways. Finally, intergenerational correlations could reflect “shared-determinants” effects, whereby parents and children share characteristics that influence the probability of being lone parent and/or in low income, in which case the intergenerational correlation would be (statistically) “spurious”, rather than causal, meaning that policies targeted on the family of origin would have limited effect.4

The U.S. research has typically focussed on mothers’ and daughters’ receipt of Aid to Families with Dependent Children (AFDC), or the subsequent Temporary Assistance for Needy Families (TANF). Because these programs are specifically aimed at providing assistance to lone parent families, the problem is a little different from the one addressed here. That is, since women must essentially be in one of the outcome states (lone motherhood) to be in the other (on welfare), the latter outcome is effectively conditional on the first, although the first does not guarantee the second. In contrast, welfare in Canada (“Social Assistance”) is generally available to individuals of all family types – unattached persons and married couples along with lone mothers – so while the outcomes may be statistically linked, they are independent to some degree. This is an important difference which carries over to the particular choice of statistical model employed in our analysis of low income status (rather than welfare receipt) as discussed below.

An, Haveman and Wolfe (1993), using 1986 to 1987 data from the Panel Study of Income Dynamics (PSID) and a bivariate probit model, find that premarital childbearing was more likely in daughters whose mothers had low education and were themselves participants in AFDC. Ratcliffe (2002), however, using 1968 to 1991 data from the PSID and a multinomial logit model, finds an association between mothers’ and daughters’ participation in the AFDC only among blacks, as well as no strong evidence of a relationship between mothers’ participation in AFDC and daughters’ premarital childbearing.

Gottschalk (1996), again using the PSID, takes into account unobserved heterogeneity to conclude that correlated unobservables are important in explaining the intergenerational correlation

4 See Gottschalk’s [1996] development of this idea in particular. See also Beaulieu, Duclos, Fortin and Rouleau (2001) for a discussion of these effects regarding the intergeneration reliance on social assistance in Québec.
in welfare participation among blacks but not among non-blacks, although the correlation across generations is significant for blacks as well as whites even after taking account of those effects.

In Canada, Corak and Heisz (1999), Corak (2001), and Fortin and Lefebvre (1998) provide evidence on the intergenerational correlation of incomes (or earnings) in general, but little research has been done on the transmission of poverty status *per se* or the inheritance of lone motherhood. This is principally due to the lack of data suitable to the task – a longitudinal dataset of sufficient breadth and length which also permits individuals to be linked to their families of origin and then followed as over time as adults. The closest work to ours is Beaulieu, Duclos, Fortin and Rouleau (2001), who look at the inter-generational correlation of participation in Quebec’s social assistance program to find important background effects, especially for periods spent on social assistance during the early stages of childhood (age 7-9) and late adolescence (16-17) – results which turn out to be quite similar to what we report below.

One set of influences that has not been specifically addressed in any of these studies is the neighbourhood in which the child grew up. In what has become a rather large literature on neighbourhood effects on various other outcomes, Oreopoulos (2002) represents a recent Canadian study which provides a review of the general neighbourhood literature and then looks at neighbourhood effects – characterised principally by local poverty rates and the percentage of lone mother families – on young men’s subsequent earnings, finding small effects. Akerlof and Kranton (2000) focus on how certain neighbourhood effects related to identity or role model effects are likely to operate on individuals’ outcomes.

While many neighbourhood effects could be important, in this paper we consider what are generally considered the most important indicators while also maintaining a symmetry with the outcome variables focused on to include just two: the percentage of families in low income, and the percentage of lone mother families. These could embody a range of effects pertaining to the three broad categories discussed above: human capital investments, preferences or values, and shared determinants.

We thus focus on the intergenerational transmission of lone mother and low income status in terms of the effects of the individual’s lone mother and low income family and neighbourhood experiences in childhood. This paper does not attempt to disentangle the separate influences of the family or neighbourhood effects into their component parts (human capital, preferences and shared
determinants). Its focus instead is to provide estimates of whether the overall effects – and which of these in terms of the low income and lone mother influences – are significant or important.

In summary, this paper fits into the existing literature by i) applying the general approach developed in U.S. work on the intergenerational transmission of lone mother and welfare participation to the lone-mother-low-income problem and thus providing the first Canadian evidence of which we are aware on these processes, ii) adding neighbourhood effects to the models, and iii) otherwise exploiting the longitudinal nature and massive size of the LAD data in a number of ways, including tracking childhood outcomes from infancy through the late teen years and differentiating the effects of each of the family and neighbourhood influences over the different periods of the child’s development.

III. THE DATA AND MODEL

III.1 The Longitudinal Administrative Databank (LAD)

The LAD is a 20 percent representative sample of Canadian tax filers constructed from Canadian Revenue Agency tax files that follows individuals over time and matches them into family units on an annual basis, thus providing individual and family-level information on incomes, taxes, and basic demographic characteristics in a dynamic framework. The first year of data is 1982 and the file ran through 2000 when this project was undertaken.

Individuals are selected into the LAD by a random number generator and are then linked over time for all the years they file a tax form. The LAD’s coverage of the adult population is very good since, unlike some other countries (e.g., the U.S.) the rate of tax filing in Canada is very high: upper income Canadians are required to file, while lower income individuals have strong incentives do so in order to recover income tax and other payroll tax deductions made throughout the year and to receive various tax credits. The full set of annual files from which the LAD is constructed are estimated to cover 95-97 percent of the target adult population in any given year.

The LAD has a number of characteristics that make it well suited to this study. First, the LAD allows us to link individuals who file taxes while still living at home (or who use their parents’ address) – the majority of the population – to be linked to their family of origin. We can then follow

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5 Individuals may be included for some years but not others in which they do not file a tax form. This boosts the representativeness of the file in a longitudinal framework relative to sampling regimes where a missed year results in permanent attrition from the sample.
the individual’s family and neighbourhood circumstances back in time through the mother, thus permitting us to identify the family and background variables used in the analysis.

Second, the extended period covered by the LAD allows us to look at daughters’ characteristics in childhood as far back as the age of one, and then forward as an adult to as late as 37 years of age, although individuals’ records are unbalanced in this regard. That is, individuals who are generally older over the period covered by the LAD are followed for more years as adults but are not tracked as far back in their childhood, while younger cohorts are not followed as far into adulthood but are tracked back to younger ages in childhood.6

Third, because the LAD is based on tax data, it contains detailed and reliable income information, thus permitting a precise analysis of low income status.7 Similarly, the inclusion of the person’s marital status and presence of children on the file allows us track the type of family they were living in as a child and the evolution of the individuals’ own family status, in particular lone parenthood status, as an adult.

Finally, because the LAD contains a very large number of observations (nearly six million in 2000), we are able to construct large samples of young women and identify the parameters attached to the effects of their early low income and lone mother experiences that interest us here on their own outcomes in this respect. The large number of observations on the full LAD also allows us to construct the neighbourhood characteristic variables on an annual basis with a high degree of accuracy.

III.2 The Samples and the Variable Definitions

Two samples are constructed. In sample A, we identify all woman aged 20 or above who appear in any year of the LAD and who previously filed a tax return while living at home with their parents when aged 14 to 19.8 Once this connection is made, we identify the family and neighbourhood characteristics (low income, lone mother family) used in the analysis. For these, we

6 The treatment of the samples in this regard is explained further below.

7 Statistics Canada now generally attempts to draw income data for all its surveys from individuals’ tax files, at least partly because they are deemed to be an accurate source of this information.

8 This includes individuals who simply used their parent’s address for their tax return even if they were actually living away from home at the time. It is, in fact, address matches such as these (plus other information derived from individuals’ tax forms such as name, sex, and age) by which individuals are matched into family units in the LAD.
use the daughter’s information when she was aged 16 years old if it is available, in order to have as common a year of reference as possible. If that year’s information is not available, we use the next closest age. Our first set of models is estimated with these samples.\(^9\)

Sample B includes all those daughters from sample A for whom the mother (not necessarily biological) also appears in the LAD. This link is made through the family identifier on the file, and requires that the mother is also a tax filer or imputed by a spouse with a trackable identifier – her SIN. Because the LAD is a 20 percent random sample of Canadian tax filers and because of the high rate of tax filing, this daughter-mother link was successfully made in almost exactly 20 percent of the cases. We then tracked the mother back in time to identify through her the daughter’s family and neighbourhood characteristics as far back as when the daughter was one year old, thus allowing us to compare the effects of these influences at different points in the daughter’s childhood on her own subsequent lone parent and low income status.\(^10\) Our second set of models are estimated with these samples.

The unit of analysis for sets of models is person-years, with one observation for each year beginning when the daughter reaches 20 years of age. Sample A includes the background family and neighbourhood information only for a single year centred on age 16 (as described above), and has 2,474,300 person-year observations, including repeated observations for each individual according to how many years they are included in the LAD from age 20 onward. Sample B also follows women from the age of 20, but includes the extra background information from the daughter’s earlier childhood years derived from the mother matches, and has 517,220 person-year observations.

\(^9\) Approximately 65 percent of the young women captured in the LAD are linked with their family of origin before the age of 20 in this way. This lack of full coverage presents the possibility of selection bias, but there are a number of reasons to believe this bias is not likely to be too large. First, there is no reason to expect, \textit{ex ante}, such a bias, since filing or not filing at home would not seem to be obviously related to any particular individual or family situation. Second, there are only small differences in the current (i.e., when adult) lone mother and low income status of matched versus non-matched women. Third, the model results were very similar when they were estimated with those who filed at each age (16 or under, 17 or under, 18 or under, etc.), which is at least consistent with filing not being related to the underlying dynamics. Finally, other published work using similar tax-based data to estimate roughly similar intergenerational processes, including Corak and Heisz (1999), Corak (2001), and Oreopoulos [2002], found no evidence of any such selection bias.

\(^10\) Care is taken to ensure that there is a continuing mother-daughter match that likely corresponds to the daughter in question as the mother is tracked back in time. In particular, there must be a child of the appropriate age in the mother’s family in the earlier years. If not, the match is broken at that point. This process allows us to censor records at the point it seems likely that the child was not living with the woman in question, such as can occur with reconstituted families where the identified “mother” is not the daughter’s own parent. We follow mothers on the grounds that children remain with their mothers in the vast majority of cases, no matter how family circumstances evolve over time.
The background variables are thus fixed for the women’s adult years which enter our models, while her own low income and lone mother circumstances of course vary – indeed to a considerable degree, as shown below. Standard errors are corrected for repeated observations on the same individuals.11

The key family characteristic variables used as both explanatory variables (the daughter’s family situation when at home) and as outcome variables (the daughter’s own situation later in life) are low income status and being in a lone mother family. The latter is defined in a straight-forward manner, while we define “low-income” as situations where adjusted family “market income” is less than one half the Canadian median for the year in the question. The square root rule is used to adjust family income for the number of individuals in the household. Both the square root rule and the one-half median cut-off represent commonly used conventions.12 As for the use of market income rather than total income, this was necessitated by the absence of certain important government transfers (social assistance in particular) from the earlier years of the LAD. The measure represents however a good proxy of economic well-being and is suitable for our purposes.13

The neighbourhood characteristics used as explanatory variables are analogous to the family variables: the percentage of families in low income in the neighbourhood, and the percentage of lone parent families. Neighbourhoods are defined by Canada Post’s Forward Sortation Areas (FSA), which correspond to the first three characters of a postal code and represent a set of well-defined and rather stable areas. The population count in each FSA varies, depending notably on whether the area

11 Because post-secondary students typically have lower earnings than non-students, precisely because they are in school rather than representing any more significant hardship, individuals were excluded from the analysis for the years they were identified as students, but then included when this was no longer the case. Student status is identified using an algorithm developed by the authors based on various tax deductions available to post-secondary students reported on the LAD. This exclusion did not, however, affect the findings in any particularly significant manner.

12 Our findings are robust to other low income measures, including Statistics Canada’s established Low Income Cut-Offs. The LAD uses the census definition of families, thus including at most two generations and precluding extended family households.

13 The market income measure includes wage and salary income, net self-employment income, investment income, and virtually all other private sources of income. Tests using total post-government income measures (i.e., including government transfers and taking taxes and tax credits into account) for the years this information is available on the LAD (from 1992) generated results very similar to those obtained with the low market income measure, and the correlation of low income status by the two measures is very high. Other work by the authors using the LAD and various low income measures consistently find little difference in results according to the precise definition used.
is rural or urban. There are 1,610 FSAs in total.\textsuperscript{14} Estimating the models using neighbourhood information based on census tracts, smaller areas of about 5,000 families on average, yielded very similar results to those using FSAs.\textsuperscript{15}

The models also include a number of control variables. One is the woman’s age. After experimenting with a detailed set of one-year dummy variables running from 20 to 37 (the age over which women are included in the models we estimate), we restricted these to two dummy variables representing ages 26 to 30 and greater than 30 (20-24 is the omitted baseline group) as this made estimation easier and had no effect on the key variables of interest. Calendar year is also included in the form of a set of dummy variables. Other control variables include current province of residence and area size of residence (from large urban area to rural).

The sample B models also include a set of variables to deal with the problem that we have information on commensurately fewer women’s backgrounds the further back in childhood we go (age 10-14, 5-9, 0-4), as discussed above. Rather than restricting the analysis to only daughters for whom the information is available for all age periods, we set the variables to zero where it is missing and set a corresponding “missing data” dummy variable (one for each age interval) to one to control for the missing effects. Because this is a simple, if potentially effective (Card (****)) method of dealing with such missing information in situations like these, we also estimated various models restricted to individuals for whom the information for all the background years included in the models was available, and found this to have no significant impact on our main findings. We therefore use the samples which include those with missing information and add the associated controls because they provide more observations, they provide estimates of the relevant effects over a more extended period of time and more cohorts of young women, and they minimise any selection bias that might be related to the consistency of mothers’ tax filing over extended periods.

One last control variable is included in the low income models: an indicator of whether or not the young woman was still (currently) living at home. This variable was added in order to control for situations where the woman’s current economic status was a function of her not having yet struck out on her own, rather than her own independent situation. A more complete treatment

\textsuperscript{14} See Statistics Canada (2003) for more details.

\textsuperscript{15} Since we already weight records for the repeated observation on given individuals, we did not also weight for the repeated observations on individuals from the same neighbourhoods, since the former would have the greater effect and it is not possible to do both.
would include modelling that outcome along with low income status, but such an approach is beyond the scope of this paper. An alternative simple approach is to not include such a control variable and allow the woman’s low income status to effectively be estimated in a reduced form manner where it might operate to at least some degree through her choice of living arrangements. This was done, but the main findings were again not appreciably altered by the treatment of this control variable, indicating that the issue is not important in a practical sense.

Finally, it is worth noting that a quirk of the LAD means that it is not possible to differentiate lone fatherhood from lone motherhood in sample A. This is because we are essentially using the information on the “daughter’s” record regarding her family type, and in the LAD there is no way to determine the gender of the lone parent for “filing children” who are in such families. In 2001, however, lone parent families were headed by the father in only 19 percent of all cases (see above) and this proportion was progressively lower in earlier years. The upshot of this lacking information is that the models estimated with sample A capture the effects of being in a lone parent family (be it headed by a mother or a father), while the models estimated with our sample B (where we follow the woman’s mother back in time) capture the effects of being in a lone mother family. We have, however, estimated the more restricted (“sample A”) model with sample B, and the results are largely similar.

### III.3 The Models

The models we use to explore the intergenerational dynamics of lone motherhood and low income status extend from the established literature, and thus consist of a discrete choice probability framework where the daughter’s outcomes are defined as whether or not she is a lone mother or in low income status (or both) in year $t$, and these are taken to be a function of the vector of control variables, $X_{it}$, mentioned above (age, calendar year, area size of residence, province, etc.), plus the childhood family and neighbourhood effects. Specifically:

$$
\text{Prob}(y_{it}) = X_{it}\theta + \beta_{\text{childhood}}i + \epsilon_{it}
$$

The definitions of the childhood characteristics variables vary with the sample used. In sample A, where we make no attempt to track mothers and use only one year of information, we define the following variables, all corresponding to the situation when the young woman was 16 years of age (or as close to 16 as possible):
mthr_lp: takes a value of one if the daughter’s family was a lone parent family, zero otherwise.

mthr_li: takes a value of one if the daughter’s family was in low income, zero otherwise.

nbrd_lp: the proportion of lone parent families in the daughter’s neighbourhood.

nbrd_li: the proportion of low income families in the daughter’s neighbourhood.

In sample B, where we track the daughter back in time (through their mothers), the same kind of variables are defined, but with two differences: a separate variable is defined for each period of the daughter’s childhood (1 to 4 years old, 5 to 9 years old, 10 to 14 years old and 15 to 19 years old), and the measures represent the average situation over each of these intervals. We thus define:

mthr_lp: the percentage of years, over those for which the information is available, the daughter’s mother was a lone parent during each of four above-defined daughter’s age periods.

mthr_li: the percentage of years, over those for which the information is available, the daughter’s family was in low income during each of the same four above-defined age period.

nbrd_lp: the average proportion of lone parent families in the daughter’s neighbourhood for each of the four above-defined daughter’s age period.

nbrd_li: the average proportion of low income families in the daughter’s neighbourhood for each of the four above-defined daughter’s age period.

We use a number of specific econometric specifications to estimate the relevant intergenerational dynamics. First, we use simple logistic (logit) regressions to model separately the probability of adult daughters being i) a lone mother, ii) in low income or iii) a lone mother in low income. Such results, however instructive regarding lone motherhood and the status of low income, reduce the underlying set of issues to a set of simple bivariate comparisons.

Following Ratcliffe (2002), our second approach is to use a multinomial logit model which allows us to consider a broader set of combinations of marital status and low income status. The daughter can be, at age t, 1) a lone mother in low income, 2) a lone mother not in low income, 3) single without children, in low income, 4) married or in a common-law union in low income or 5) married, in a common-law union, or single and not in low income. The fifth outcome is the baseline outcome because it represents a situation where the woman is neither a lone mother nor in low income.
Our third, and preferred approach is to employ a bivariate probit model which permits the joint modeling of the status of low income and lone motherhood status while allowing for a correlation in the error term between these two events, while also relaxing the Independence from Irrelevant Alternatives (IIA) assumption of the standard logit model.

IV. EMPIRICAL FINDINGS

IV.1 Simple Rates of Lone Motherhood and Low Income Status

Table 1 provides some descriptive statistics regarding lone motherhood and low income status by age for the young woman in our two samples as we track them over time. As could be expected, the percentage of women who are lone mothers increases with age, from 3.4 percent at age 20 to 11.7 percent at age 37 (sample A), while the percentage in low income decreases, from 18.8 percent to 9.6 percent. The latter tendency reflects the fact that women’s earnings tend to grow with age, while they also move into couple relationships where the spouse’s earnings (which also increase with age) boost family income and economies of scale in living arrangements are realised. The lone mother pattern reflects women’s increased probability of having at least one child as they grow older, while some of them find themselves with no partner either because of divorce or because they were never married.

The net effect of these two opposing tendencies is the percentage of young women who are both lone mothers and in low income. The rates are overall quite low but first rise sharply and then gradually decline, from 3.1 percent at age 20, to a maximum of 4.9 percent at age 23, tapering off to 4.1 percent at age 37.

The table also shows that younger lone mothers are more often in low income than older lone mothers. At 20 years old, more than 9 lone mothers out of 10 are in low income. This proportion slowly decreases, but even at 31 years old, the majority of lone mothers are in low income. At 37 years old, the oldest women can be in our samples, more than one woman out of three is in low-income. These simple statistics underline the importance of investigating lone motherhood and low income status jointly, especially at younger ages.

IV.2 Regression Results: Sample A (Effects at Age 16)

The first set of results, for Sample A, where the family and neighbourhood characteristics at age 16 are included, are shown in Table 2. These identify the principal themes which hold through
the more detailed specifications (i.e., using sample B). The upper panel of results are for the lone mother outcome, the bottom panel for the low income outcome. The comparable simple logit model results are shown in Table 4. Since the latter results are very similar to the bivariate probit results, we only discuss the former, which represent our preferred specification on theoretical-econometric grounds.

The first result of note is that the family and neighbourhood coefficient estimates are all statistically significant at conventional significance levels, but with such large samples this is not surprising, and we focus instead on the estimated magnitudes of the effects to identify the patterns of importance. A more interesting general result is the $\rho$ coefficient, whose estimated value of .700 points to a significant correlation of the two outcomes and the appropriateness of using the bivariate probit model to model the lone mother and low income outcomes in a joint fashion.

Regarding the lone mother outcome, the findings indicate that being in a lone mother family at age 16 (or thereabouts) is strongly correlated with the probability that a young women is later herself a lone mother. The estimated effect is a 3.9 percentage point increase in the probability of being a lone mother in any given year, or a 76.1 percent increase relative to the baseline rate of 5.1 percent (i.e., where all regressors are set to zero). Having been in a low income family has a strong additional effect of 3.0 percentage points, or an increase in the probability of being a lone mother of 58.7 percent in relative terms.

Neighbourhood effects also matter, but the results are not nearly so symmetric. Even after taking account of the young women’s own family situation – of being in a low income or lone mother family – being in a neighbourhood with a higher percentage of lone mothers has a strong influence on whether or not the woman goes on to become a lone mother. More specifically, doubling the percentage of lone mothers in the neighbourhood in which the woman was living at age 16 (from the national mean of 9.3 percent to twice that level) raises the probability of being a lone mother 4.3 percentage points, an 83.8 percent increase. Conversely, being in a low income neighbourhood (defined in the same manner of doubling the rate from the national average of 17.8 percent to twice that level) has no effect, the point estimate actually indicating a slightly negative effect.

Turning now to the low income outcome, we see that the family effects again matter, but the effect of the low income measure is now much stronger than the lone mother indicator – increasing the low income probability by 106.4 versus 38 percent. Furthermore, the neighbourhood effects are
now both quite weak, with the same doubling of rates as before now increasing the probability almost negligibly.

These results are significant for a number of reasons. First, a young woman’s family background appears to in fact be a very important determinant of her own lone mother and low income status, while the fact that having been in a lone mother family has an independent effect in addition to having been in low income points to the special problems of lone mother families.

Second, the fact that the only neighbourhood effect that matters is being in a lone mother neighbourhood for the lone mother outcome might suggest a role model effect. If it was purely neighbourhood resources that mattered (schools, activities, etc.), it would seem that the low income neighbourhood measure should be most important. If alternatively, neighbourhood effects risk picking up unobserved family effects, again it would seem these would be most likely captured by the low income neighbourhood background measure. But the low income background variable is not significant, while the lone mother measure is. So either the lone mother measure is picking up other unobserved/omitted effects in a special way, or it is in fact the lone mother aspect of the neighbourhood that matters. This is an intriguing result with potentially interesting policy implications.

One limitation of the bivariate probit model is that only binary outcomes can be considered, and the bivariate model inherently places a certain structure on the nature of these relationships. To allow various explicit combinations of low income and lone motherhood (or more generally marital status) outcomes to be modelled jointly, we use multinomial logit regressions. The results for Sample A are shown in Table 6.16 The possible outcomes for the daughter, at age t, are 1) lone mother in low income, 2) lone mother not in low income, 3) single (unmarried) without children in low income, 4) married or in a common-law union (with or without children) in low income or 5) married, in a common-law union, or single and not in low income (the baseline outcome).

The results of most interest are for the outcome of being a lone mother in low-income against the baseline outcome of being married, in a common-law union or single and not in low-income. The results here are consistent with those of the bivariate probit model and are particularly

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16 Because of the relative complexity nature of the multinomial regression model, the associated predicted probabilities do not have as intuitive an interpretation as those from the simple logistic or bivariate probit models. Therefore, following Ratcliffe (2002), only coefficient estimates and associated p-values are presented and discussed.
conclusive with regard to the family effects. Being in a lone parent and/or a low income family at age 16 is strongly and positively associated with a higher probability of the young woman then being a lone mother living in low income as opposed to being either attached or single without children not in low income. The neighbourhood effects seem once again to play some role, especially the proportion of lone parent families in the neighbourhood where the woman lived at age 16.

IV.3 Regression Results: Sample B (Effects by Age)

Our findings using sample B, which include the family and neighbourhood effects over the different childhood age intervals, are shown in Table 3. The results obtained with Sample A generally hold in that the family effects generally matter to both outcomes, while the only neighbourhood effect of real strength is the lone mother effect on the lone mother outcome. But the timing of the effects is the focus here and the results show some interesting patterns, while some additional, more subtle neighbourhood effects also become evident.

Perhaps the most striking result is the impact of exposure to low income and lone mother status in the family at the dawn of adulthood. Daughters who were in a lone mother family every year between the age of 15 to 19 have a predicted probability of themselves being lone mothers which is 63 percent larger than those who had no experience of lone motherhood during the same period. The impact of exposure to lone mother status is smaller at earlier points in the woman’s childhood, but still strong for ages 0 to 4 and 5 to 9 (around 30 percent). It is only over the age 10-14 interval that the effect is really quite small (6.5 percent).

Similarly, daughters whose family was in low income every year they were 15 to 19 have a predicted probability of themselves being lone mothers that is 89 percent greater than those who did not experience low income during the same period. Again, the impact of exposure to low income status diminishes as we go back in time to consider earlier stages of childhood but here the pattern is a little different, with the 10-14 interval remaining strong and the earlier periods being least important.

Continuing with the family effects, daughters who experienced lone motherhood in the late teen years are 33 percent more likely to be lone mothers at adulthood than daughters who did not (the probability is raised from 3.2% to 4.8%). This intergenerational phenomenon is however not as clear if we consider earlier stages of the daughter’s childhood. There seems to be practically no
effect if we consider early teens (10 to 14 years old), a somewhat more effect for mid-childhood (23 percent), and again little effect for the youngest group.

Similarly, daughters whose family was in low income when they were between 15 and 19 years old have, on average, a probability of being themselves in low income that is more than twice that (an increase of 116 percent) of daughters whose family was never in low income during that period. The amplitude of the effect is more than cut in half, however, if the family exposure to low income status occurred at early adolescence (49.6 percent). Growing up in a low income family before the age of 10 still matters, but less, with the effects being 24.2 and 19.8 percent, respectively, for the 5 to 9 and 0 to 4 age groups.

Compared to these family effects, the evidence of neighbourhood effects clearly is again not as convincing overall. The proportion of low income families in the daughter’s neighbourhood has virtually no impact on her likelihood to either be in low income or a lone mother at adulthood. The only exceptions to this rule are smallish effects on the low income outcome if the neighbourhood effects were experienced either early (age 0 to 4) or late(15 to 19).

The proportion of lone parent families in the neighbourhood, however, still matters, especially for the lone mother outcome, and the timing of these effects is interesting. Daughters who, when aged 15 to 19 years old, lived in a neighbourhood whose proportion of lone parent families was double that of the sample average see their probability of being a lone mother being raised from a baseline of 2.6 percent to 4.6 percent, or an increase of 74 percent in relative terms. The effect is somewhat smaller for the 5 to 9 years old period (42.7 percent), and much smaller than this for the other intervals.

Unlike what was found with sample A, the proportion of lone parent families in the neighbourhood also has a moderate effect on the daughter’s probability of later being in low income, especially – again – if that neighbourhood exposure cam in the late teen or mid-childhood years.

V. CONCLUSION

In this paper, we have used the Longitudinal Administrative Databank (LAD), a 20 percent representative sample of Canadian tax filers, to study the impact of growing up in a lone parent and/or low income family on a young woman’s probability of herself being a lone parent and/or being in low income status one generation later. Neighbourhood characteristics when the woman was a child are also incorporated into the model.
The results show evidence of a strong intergenerational transmission of lone motherhood and low income status. Growing up in a low income family and/or in a lone parent family increases, often dramatically, a young woman’s probability of herself being in low income and/or a lone mother. The findings also identify the effects of neighbourhood characteristics, the proportion of lone parent families being most important and its effect being greatest with respect to the lone parent outcome.

These results are generally quite robust across different models and samples. At the family level, being in a low income family matters more to the daughter’s outcomes than being in a lone parent family, while the opposite is true at the neighbourhood level, as the proportion of lone parent families matters more than the proportion of low income families. These various influences are, furthermore, while relatively consistent over different age ranges, seen to be strongest in the early teen years.

For policy makers, these results have several important implications. At a minimal, descriptive level, the results point to a means of identifying young women at greater risk of experiencing lone parenthood and low income in early adulthood and thus of targeting associated policies.

In addition, if subsequent research finds that this intergenerational transmission is indeed causal, then policies focussed on families with children could not only have immediate impacts on those families at that point in time (e.g., alleviating poverty), but also on the children of those families in later years as they become adults and form their own families. The question here is whether or not we can assess in a satisfactory fashion the causality of the relationship.

One way to do this, following Gottschalk (1996), would be to add to the set of explanatory variables characteristics of the mother when the daughter is an adult. If the intergenerational transmission mechanism identified here merely reflects correlated unobservables, and is therefore spurious rather than causal, those variables will have a similar effect on the daughters’ outcomes as those which capture the family situation while the girl is living at home. Unfortunately, conducting this exercise with the lone motherhood variable is not a very meaningful exercise, since family status depends in a definitional manner on the presence or absence of the daughter, and even low income status is subject to similar important caveats, since it depends on family size and composition.

However, as the theory on human capital and imitation effects reviewed earlier discusses, there are strong reasons to think that the relationships identified here are indeed causal, at least to
some degree. In particular, the neighbourhood in which children grow up can be seen as purely exogenous to the child; in other words, young children do not choose the neighbourhood in which they grow up. As such, the effects of neighbourhood characteristics after controlling for family characteristics could probably safely be considered as being causal.

We thus conclude that the results presented here point to an interesting set of empirical relationships that have not previously been identified that at least some of these are likely to in fact be causal rather than simple empirical correlations. But we fully acknowledge that other work is required to validate the exogeneity issues in particular, while the work could undoubtedly be taken in any number of other directions as well.

One particular line of future research would be to add social assistance (SA) participation to the framework, along the lines of the work carried out by Beaulieu, Duclos, Fortin and Rouleau (2001) on SA participation in Quebec. In addition to being more directly comparable to the existing U.S. literature, it would address the important issue of lone motherhood and low income in light of one of the most important and controversial social programs in Canada.
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