An Econometric Analysis of Intergenerational Reliance on Social Assistance *

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Abstract

This paper examines the intergenerational transmission of participation in Québec's social assistance program. The analysis takes into account two sources of intergenerational transmission: one that is due to a causal link between parents' and children's participation and one that is due to a correlation between individual or environment-specific characteristics across generations. Our data come from the records of Québec's *Ministère de la Solidarité Sociale* and cover 17,203 young people who were 18 years old in 1990 and whose parents were recipients of social assistance during at least one month between 1979 and 1990. Our results reveal that, on average, a one-month increase in the parental participation during the youth's pre-adult years (age 7–17) raises the youth's participation by about 0.15 month during early adulthood (age 18–21). Moreover, this impact is stronger during the early stages of childhood (age 7–9) and late adolescence (age 16–17).

Keywords: Dynamics of social assistance participation, Intergenerational correlation, Canadian welfare programmes. JEL #: C34, C35, H53, I38.

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1 Introduction

In Canada, provincial social assistance (SA) benefits are paid to those whose income is deemed insufficient, whether permanently or temporarily. Thus the *SA* programs have the direct effect of reducing the incidence of poverty and of improving the financial situation of families with limited resources. These programs have nonetheless been criticized as generating undesirable effects on recipients and on their environment.[1] One such effect that is frequently mentioned is the transmission of past program participation by parents into an increased probability of future participation by their children. To our knowledge, however, this effect has not been investigated or documented in Canada, though Corak *et al.* (2000) report evidence on intergenerational receipts of unemployment insurance (UI) benefits.[2] Our paper sheds light on this issue by estimating the magnitude and the nature of intergenerational reliance on *SA* in the province of Québec.

An observed correlation between the participation profiles of parents and their children may originate from two distinct sources: there may be a causal link between parents' participation and that of their children, and there may be intergenerational correlation between individual or environment-specific characteristics that affect the propensity to participate.

Two main reasons may explain the presence of a causal link. First, the participation of parents may reduce the family stigma felt by their children when they use the *SA* program. This effect may therefore reduce their reluctance to rely on the program (an imitation effect). Second, children may learn how to use the program while living with parents receiving *SA* benefits (a learning effect). They may therefore face lower participation costs when they grow up, which encourages them to use the program. Through these links, parental participation in the *SA* system, at a time when the youth has not yet reached adulthood, can increase the probability that the latter will participate in the future.

It is also recognized that certain observed and non-observed characteristics tend to be correlated across parents and their children (*e.g.*, the level of education, the motivation to work, the neighborhood of residence), and that these characteristics can affect the degree of reliance on *SA* (a shared-determinants effect). For instance, both the parents and the children may perceive work as involving large nonpecuniary disutilities, for social or family-specific reasons. In this case, the exclusion of young adults from the labor market, and consequently their reliance on *SA*, spring from values and perceptions commonly shared by the parents and the children, and is not directly caused by the parents' participation in the *SA* program [*e.g.*, An *et al.* (1993) and Duncan *et al.* (1988)].

[1] For a review of the issues and of the American evidence, see Moffitt (1992).

[2] Also Corak and Heisz (1999) find some evidence of intergenerational earnings and income mobility in Canada.

For public policy reasons, it is important to identify correctly not only the magnitude, but also the transmission mechanism of intergenerational reliance on *SA*. In the specific case in which the child replicates the *SA* behavior exhibited by the parent due to an imitation effect or a learning effect, there will be a causal link between parents' and children's participation in *SA*. In such a case, policies which impact on the participation of parents will also affect that of their children. This implies that any cost-benefit analysis of these programs should account not only for their impact on the current generation, but also for their impact on future generations. For instance, in the presence of a causal link, implementing more stringent eligibility criteria for the parents may result in a reduced participation rate for their children. However, even when *SA* participation is correlated across generations, the impact of tightening parental eligibility criteria will have little or no influence on youth participation to the extent that the correlation of parents' and children's behavior is spurious. This will be the case when it reflects the sharing of values and attitudes generated by their common general living environment – the shared-determinants effect. In this case, a more effective policy to decrease future reliance on *SA* would be to improve the children's environment and those socio-economic characteristics that are determinants of *SA* participation.

In Section 2 we review in some details the principal factors that distinguish between causality and shared-determinants effects. We also illustrate how it is possible to capture this distinction statistically. Section 3 presents the econometric model. The sample is described in Section 4. The results of the econometric estimation are presented in Section 5. We conclude our analysis in Section 6 by proposing *inter alia* some avenues for further research.

2 Causality versus shared determinants

As discussed above, the intergenerational transmission of reliance on *SA* can take two distinct forms, depending on the nature of the transmission mechanism. Observed correlations between participation rates may result from a causal link between the behavior of parents and children, or may be attributable to the correlation across generations of observed and unobserved determinants of *SA* participation.

The notion of a causal link refers to a natural replication of the parental *SA* model by the children. The mere fact that the parents are on *SA* during some period of the youth's pre-adult life may provide an incentive for him to participate in the program. Several factors can generate this causal link.

First, parental participation in the program can change the youth's preferences by reducing the family stigma associated with the status of *SA* recipient. The youth may thus manifest an increased motivation to become a *SA* recipient simply because his parents were *SA* recipients during his pre-adult life, namely, simply to imitate the family behavior [An *et al.* (1993), Corcoran *et. al.* (1992), Pepper (1995), Levine and Zimmerman (1996), Gottschalk (1996)].

Second, parental welfare participation may lower the *SA* participation costs of offspring. Parents already well informed of the procedures of obtaining *SA* can show their child how to use the system more efficiently. Unlike a child whose parents are active in the labor market and know nothing of the program's characteristics, a child of recipient parents can more readily obtain information and will have a lesser need to inform others of his wish to apply for benefits [Antel (1992), Moffitt (1992), and Gottschalk (1996)].

Third, owing to their status as *SA* recipients, recipients are less equipped to offer their offspring opportunities to develop job-search skills. Furthermore, since they have fewer interactions with the labor market, they also have less scope to provide their children with contacts to people with information and influence on that market [Duncan *et al.* (1988), Antel (1992)].

Finally, the features of Québec's *SA* system may also partly explain the existence of a causal link between parental and child participation in the program. Indeed, in the Québec *SA* system, the loss of student status entails a reduction in the benefits paid to the parents when the youth reaches 18 years. By collecting *SA* for himself, the youth can partially compensate for these foregone benefits. There is some evidence of this. In a recent survey on the Québec *SA* program, some young claimants – when asked to give their reasons for participating in *SA* – reported that they were pressured by their parents [Lancôt and Lemieux (1995)]. Youths above 18 years were especially prone to be asked by their *SA* parents to claim *SA* if they had dropped out of school.

The second basic explanation for an observed correlation between parents' and children's participation relates to the correlation of observed and unobserved *SA* participation determinants across parents and children. For instance, the correlation across parents and children of attitudes towards their socio-economic environment plays a likely role in explaining the correlation in their socio-economic decisions. Thus, if the nonpecuniary disutility of working is perceived as very large by both parents and children, voluntary non-participation in the labor market (and thus reliance on *SA*) can be viewed as a rational decision by both generations [Duncan *et al.* (1988) and Solon *et al.* (1988)]. Moreover, some personality traits, possibly difficult to observe directly, are plausibly transmitted from parents to children, and these may impact on their common willingness to receive *SA* [see Duncan *et al.* (1988), Levine and Zimmerman (1996) and Gottschalk *et al.* (1994), and especially Herrnstein and Murray (1994) for genetic evidence].

In our empirical work, we will not attempt to identify each of these numerous transmission mechanisms. However, under some assumptions, our econometric approach will allow us to isolate the general causality effect from the shared-determinants effect. To illustrate how one can identify these two effects, consider an hypothetical example in which the observed and unobserved characteristics of parents (and thus the determinants of parental reliance on *SA*) are the same before and after their children have reached adulthood. For simplicity, further suppose that the observed pre- and post-adult periods for the children are of identical length, and that the *SA* participation rate of children once they have become adults is deterministically set by their characteristics (which are correlated with those of their parents) and by the pre-adult participation rate of their parents. This latter assumption, which focuses on the timing of parents' and children' behavior, is essential to the identification of the causality effect in the context of this model. Table 1 at the end of the paper shows the participation rates of four such hypothetical pairs of parents and children. These participation rates are defined as the proportion of time spent on average on *SA* by the parent or by the adult child.

Comparing these four pairs enables us to show the causal effect of parental participation on the participation of children as well as the effect of the correlation of *SA* determinants across generations. Indeed, the data of Table 1 was built in such a way that the child's adult participation rate is the sum of half the parent's overall average participation rate (the effect of the intergenerational correlation of *SA* determinants) and of half the parent's participation rate prior to the child's adulthood (the causal effect). Comparing pair 1 with pair 2, we note that the participation of parents is on average the same for the whole observed period of time (40% + 60% versus 60% + 40%). Thus, the effect of parents' characteristics on their own *SA* participation is roughly the same. If intergenerational welfare transmission only took place through the transmission of the parents' characteristics, we would expect the participation rate of children 1 and 2 to be the same. This is not the case, however, since the child whose parent had lower pre-adult participation (pair 1) also has lower welfare participation. This would indicate the presence of a causal effect in the intergenerational transmission of *SA* dependence, since, for equal overall parental propensity to rely in welfare, greater participation prior to adulthood leads to greater child participation once he has become adult. Comparing pairs 3 and 4 leads to the same observation: average parental participation is the same in both cases, but when greater parental participation is observed before the child's adulthood, this is causally transmitted into greater child *SA* participation once child adulthood is reached. Statistically, this introduces a greater correlation between children's *SA* participation and pre-adult parental participation than between children's *SA* participation and post-adult parental participation. We shall find this in the description of our own data below.

Table 1 also shows the effect of the intergenerational correlation of characteristics associated with greater *SA* participation. Parents 1 and 3 and parents 2 and 4 have identical participation rates prior to the adulthood of their children. Hence, if intergenerational *SA* transmission only took place through the causal transmission of the parents' participation prior to their child's adulthood, we would expect the participation of children 1 and 3 and of children 2 and 4 to be the same. This, however, is not the case since children whose parents experience overall lower *SA* reliance also experience lower *SA* participation. Intergenerational correlation of reliance on *SA* in Table 1 would thus be explained both by causal transmission of parents' behavior prior to their children's adulthood, and by the intergenerational transmission of determinants of *SA* participation.

3 The econometric model

The existing econometric literature on the intergenerational transmission of welfare participation uses mainly two approaches. The first one attempts to model how long it takes for an adult individual to collect welfare for the first time as a function of earlier parental welfare participation and other (possibly time-varying) covariates. This approach is usually based on event history methods [*e.g.,* McLanahan (1988), Gottschalk (1992,1996)]. In some studies, unobserved family-specific heterogeneity is controlled for by introducing future parental welfare participation as an additional covariate [*e.g.,* Gottschalk (1996); see also Corak *et al.* (2000) for a similar analysis applied to UI]. This way of treating heterogeneity requires the assumption that earlier parental participation may affect children's participation, but that children's participation does not affect that of their parents.

The second approach aims at modelling simultaneously the sequence of welfare participation for both parents and children. Unobserved heterogeneity is introduced by allowing the presence of a correlation between the error terms reflecting the permanent unobserved propensity to participate of parents and of their children. This approach has been first used by Antel (1992) within a bivariate limited dependent model. It has also been adopted by Gottschalk (1996) within an event history model.

Our econometric model uses the second approach and therefore leads to the simultaneous estimation of the *SA* participation of parents and their adult children. More precisely, let $y_{i,t}$ be a binary variable taking the value 1 if youth i received *SA* during month $t,$ and 0 otherwise, and let $Y_{i,t}$ be a binary variable representing the corresponding information for i's parent. We denote by $\{w_{i,t}\}$ a stochastic process defined by a series of time-indexed observations on $w_{i,t}$. The expression $\{y_{i,t}\}$ thus includes the set of the observed episodes on *SA* for child i (as an adult), and $\{Y_{i,t}\}$ the set of observations on the parent before and after the child reaches adulthood. We assume that the explanatory variables for the parent, $X_{i,t}$, and for the child, $x_{i,t}$, are weakly exogenous and that they are generated by a process which is independent of the parameters of the joint distribution of $y_{i,t}$ and $Y_{i,t}$. Thus, we can focus on the estimation of the parameters of the conditional density function $f\left(\{y_{i,t}\},\{Y_{i,t}\}|\{x_{i,t}\}\{X_{i,t}\}\right).$ Since the variables for the observed episodes are binary, we assume that they are generated by a latent process based on the variables $y^*_{i,t}$ and $Y^*_{i,t}$, which in turn are defined as:

$$\begin{aligned} Y^*_{i,t} &= X_{i,t}B + U_{i,t}, \\ y^*_{i,t} &= x_{i,t}\beta + \delta Y^-_{i,t} + u_{i,t}. \end{aligned}$$

Thus, we postulate that $Y_{i,t} = 1$ if $Y_{i,t}^* \geq 0$, and $Y_{i,t} = 0$ otherwise, and analogously for $y_{i,t}$. The vector $Y_{i,t}^-$ represents the parental participation rates during the periods immediately preceding t (or preceding adulthood, depending on the specification). The vector of coefficients δ allows us to identify the causality effect described in the previous sections. We assume that the error terms are normal i.i.d. across the pairs i of parents and children, and are defined such that $U_{i,t} = V_i + \epsilon_{i,t}$ and $u_{i,t} = \pi V_i + e_{i,t}$. The term V_i captures the unobservable (permanent) characteristics shared by the parents and their children. The terms $\epsilon_{i,t}$ and $e_{i,t}$ represent the components of the error terms which are distributed independently (and identically) over time. If σ_V^2, σ_U^2, and σ_u^2 represent the variance of V_i, $U_{i,t}$ and $u_{i,t}$ respectively, the correlation (ρ) of the error terms is given by:

$$\rho \equiv \text{corr}(U_{i,t}, u_{i,t}) = \frac{\pi \sigma_V^2}{\sigma_u \sigma_U}. \tag{1}$$

Failure to account for this correlation would result in a bias in the estimate of the parameter δ because of the correlation between $Y_{i,t}^-$ and V_i. Notice that the lagged values of the variables $Y_{i,t}^*$ and $y_{i,t}^*$ do not appear in the model. Consequently, we have assumed that there is no intertemporal dependence in individual spells on *SA*. Alternatively, the retained specification may be interpreted as a reduced form of the true process generating the values of $Y_{i,t}^*$ and $y_{i,t}^*$. A comprehensive analysis of the presence of intertemporal dependence, given the possible existence of correlation between the error terms in the equations for the parent's and the children's spells, would require the estimation of a simultaneous multi-spell transition model that we leave for future research.[3] On the other hand, our sample is truncated since all parents in our administrative *SA* data have incurred at least one spell on *SA*. Our estimating approach takes that truncation into account. Furthermore, one limitation of our data is that observations on individual-specific time-varying covariates are not in general available when an individual does not receive *SA* benefits. For this reason, we will assume that covariates are not time-varying. [4] Also, as in earlier studies by Antel (1992) and Gottschalk (1996) and Corak *et al.* (2000), we assume that the causal link, if it exists, works from previous participation by the parents to subsequent participation by the children, and not *vice versa* —which seems a plausible hypothesis.

The specification of the distribution of the latent variables, $Y_{i,t}^*$ and $y_{i,t}^*$, naturally leads us to model the observed distribution of $\left\{Y_{i,t}^*\right\}$ and $\left\{y_{i,t}^*\right\}$ as bivariate probit. Since many observations are available (12 per year) for both the children and the parents, we estimate the bivariate probit model using a process of repeated observations [*e.g.*, Gouriéroux (1984)].

[3]While some studies have developed econometric techniques that handle multiple spells, for example to analyze recidivism in public programs participation [*e.g.,* Bonnal *et al.* 1997], these models do not allow for a simultaneous analysis of various transition processes.

[4]This missing variables problem also makes a multi-spell transition model difficult to estimate.

The log-likelihood function for the truncated version of the model is:

$$
\begin{aligned}
lnL(\beta,\delta,B,\rho) &= \sum_{i=1}^{I} P_{00i}\cdot ln\Phi_2(-(x_i\beta+\delta Y_i^-),-X_iB;\rho) \\
&+ \sum_{i=1}^{I} P_{01i}\cdot ln\Phi_2(-(x_i\beta+\delta Y_i^-),X_iB;\rho) + \sum_{i=1}^{I} P_{10i}\cdot ln\Phi_2(x_i\beta+\delta Y_i^-,-X_iB;\rho) \\
&+ \sum_{i=1}^{I} P_{11i}\cdot ln\Phi_2(x_i\beta+\delta Y_i^-,X_iB;\rho) - \frac{1}{T}\sum_{i=1}^{I} ln(1-[\Phi(-X_iB)]^{T-})
\end{aligned} \tag{2}
$$

where $\Phi_2(\cdot)$ and $\Phi(\cdot)$ are the cumulative density functions of a standardized bivariate and univariate normal distribution, T is the total number of months over which we simultaneously observe parents' and children's participation, $T-$ is the total number of months over which prior parental participation is calculated and P_{00i}, P_{01i}, P_{10i} and P_{11i} are proportions of the total observation period during which both youth and parent did not participate ($y_i = 0, Y_i = 0$), youth did not participate but parent did ($y_i = 0, Y_i = 1$), and so on. The last term in (2) corrects for the self-selection effect of not including in the estimation youths whose parents never claimed *SA* (a truncation effect).

The estimation of this econometric model works intuitively as follows. To capture the correlation between the unobserved characteristics of the parents and those of the child, the model implicitly evaluates the difference between the parents' observed and predicted participation rate. For example, if a parent clearly and consistently participates in *SA* at a mean rate (pre- and post-adult) that is lower than that predicted by his observable individual characteristics, there is reason to believe that the parent is influenced by specific unobserved factors which deter him from resorting to *SA*. We can thus use the statistical difference between the predicted and observed participation to verify whether some proportion π of this difference is also observed in the case of the child. In the affirmative, the parameter ρ is statistically significant, and we conclude that there exists a correlation between parents' and children's unobserved variables. If, conversely, a greater participation of the parent has no statistical effect on the behavior of the child unless this participation occurs before the child reaches the age of majority, we conclude that there exists a causal link between the child's and the parents' participation. In this case the parameter δ will be statistically significant.

Estimating this model has two main advantages. First, it involves an intuitive one-step estimation procedure that simultaneously handles the endogeneity of parents' and children's participation, the possible correlation in the unobserved heterogeneity, and a possible self-selection bias. Second, due to its incorporation of repeated observations over a given horizon, the model makes use of the complete history of children's and parents' entries on welfare and exits out of the program.

4 Data

4.1 Sample construction

The administrative files of the *Ministère de la Solidarité Sociale* provided the basic source of information for our study. We constructed our sample by extracting from those files the records of all children having reached the age of 18 years between 1982 and 1995 (whether or not they themselves were claimants at any time during these years) and whose parents had been on *SA* for at least one month between 1979 and 1995. This recovered family information on a total of 230,961 children. In certain cases, changes in family composition meant that we needed to identify a most representative parental authority for the child. We did this by defining this authority as the parent having had custody of the child for the longest period (referred to below as the "claimant parent").

The next step involved reconstructing the *SA* history of the parents for the entire period between 1979 and 1995 inclusively. Nearly 150,000 parents were identified. For the period from 1990 to 1995, the same history was constructed for those children who collected *SA* during adulthood, a total of nearly 100,000 youths. The data from both histories was then pooled and standardized (*e.g.*, we created a consistent coding for the variables across years, since various coding systems had been used between 1979 and 1995). For the purposes of this study, we retained in our working sample only those recipients who had been deemed able to work, and only those youths who were 18 years old in 1990 (and thus who were 7 in 1979 and 23 in 1995) and whose parents were born after 1930 (this ensured that all parents did not reach 65 before 1995). Finally, we also eliminated from our sample mothers who had given birth to a child before the age of 15.

Our final data set has three main advantages. First, the period for which we have information on the youths (1990–1995) allows us to draw up a relatively informative picture of their participation profile. Second, the data on parents covers a sufficiently long period to allow us to capture their participation profile for the greater part of their children's youth. Third, focussing on a single one-year cohort of young adults is likely to purge our analysis of the effect of cyclical socio-economic factors on the *SA* profile of young adults.

The final sample thus comprised data on 17,203 youths (1990–1995) merged with data on their parents (1979–1995). From this total, 9,613 youths (55.9%) never made any claim between 1990 and 1995, while 7,590 (44.1%) received *SA* for at least one month. Recall that since the data are administrative in nature, our analysis is conditional on parents having received *SA* for at least one month between 1979 and 1995. Furthermore, unlike Gottschalk (1992) for instance, we cannot deal separately with the effects of parental *SA* eligibility and parents' receipt of *SA* on the behavior of their children. Also, as mentioned above, we have little information on the individual characteristics of youths of our sample who did not incur a *SA* spell between 1990 and 1995. Apart from these caveats, however, our data compare well in scope with the American data (typically PSID) used to consider the intergenerational transmission of *SA*. Being administrative, their informative content is relatively reliable and less subject to the statistical biases (such as attrition and non-response biases) that can affect survey data. The sample size is very large, the data are monthly (in the US, the data used have often been yearly), and they cover a *SA* historical period of 17 years for parents and 6 years for their children.

4.2 Descriptive statistics

At this point it is of some interest to present some of the characteristics of the final sample used as well as descriptive correlations between the participation rates of parents and children. The main variables used for constructing the descriptive statistics are the parents' and the children's rates of *SA* participation. These rates represent the proportion of a time period during which an individual or a household receives benefits. More precisely, we define the following variables:

- **Py18** to **Py21**: the youth's annual participation rate[5] for the year between 1990 and 1994 during which he was 18, 19, 20 and 21 years old.
- **Py1821**: the youth's participation rate over the period during which he was between 18 and 21 years old.
- **Pp7** to **Pp21**: annual parental participation rate[6] for the year between 1979 and 1995 during which the youth was 7, 8, 9 and up to 21 years old.
- **Pp–**: parental participation rate during the youths' pre-adult period (7–17 years).
- **Pp+**: parental participation rate during the youths' adult period (18–21 years).
- **Pp79**: parental participation rate during the time the youth was aged between 7 and 9 (early childhood).

[5]Total number of months in which the youth participated over the year divided by twelve.

[6]Total number of months in which the parents participated over the year divided by twelve.

- **Pp1012**: parental participation rate during the time the youth was aged between 10 and 12 (late childhood).
- **Pp1315**: parental participation rate during the time the youth was aged between 13 and 15 (early adolescence).
- **Pp1617**: parental participation rate during the time the youth was aged between 16 and 17 (late adolescence).
- **Pptotal**: parental participation rate over the total period (1979–1994), *i.e.*, the period during which the youth was between 7 and 21 years old.

Table 2 provides the mean observed value for parental and youth participation rates. The parental participation rates over the early and late childhood periods of their children (Pp79 and Pp1012) are 0.41 and 0.58, and reach 0.57 and 0.52 over the children's early and late adolescence periods (Pp1315 and Pp1617). The mean participation rate of adult children (Py1821) is 0.23. Note that adult children's participation rates increase as the children become older, starting from 0.15 when they are 18 years of age and reaching 0.27 when they are 21. Part of the explanation can be found in the 1990–1993 Canadian recession that strongly affected the labor market for young workers. Table 3 presents the distribution of youths by type of family of origin, at the parent's last spell as a claimant between 1979 and 1990. The categories "single" or "childless couple" denote parents with no dependents during their last spell on *SA*. The type of family of origin is "single-parent" in 61% of the cases and "two-parent" in 39%.

Table 4 provides the coefficients of correlation between the annual participation rates of the youths and those of the parents for the youths' pre-adult (age 7–17) and adult (age 18–21) periods. For each of these two periods, results are first presented for the entire sample (17,203 observations). They are then presented for two sub-samples, according to whether the youths belong to two-parent families. The results from the whole sample, as well as those from the sub-samples, (constructed according to the type of family of origin) all reveal a positive and statistically significant correlation between the youths' participation rates and the preceding (Pp–) and subsequent (Pp+) participation rates of their parents. However, the value of the correlation coefficients between the youths' participation rates and those of the parents during the youths' pre-adult period $(Pp-)$ is greater than the corresponding results for the parents' participation rates during the youths' adult period. Moreover, equality of correlation coefficients over these two periods is statistically rejected (at the 5% level) in all cases. These results, which correspond to our expectations, are depicted in Figure 1. Table 4 shows also that the correlation coefficients pertaining to the sub-sample of youths for two-parent families are greater than those from the single-parent offspring.

It has also been suggested that the impact of parental participation in *SA* should depend on the pre-adult age of the child during which it was observed [Gottschalk (1996)]. Table 5 reports the coefficients of correlation between the participation rates of children when they are adult (18–21) and the parental participation rates when their children are in four distinct stages of their development: early childhood (7–9), late childhood (10–12), early adolescence (13–15) and late adolescence (16–17). Once again, these statistics are calculated for the entire sample and for the aforementioned sub-samples (single- and two-parent families of origin). All coefficients have a positive and statistically significant sign. As a general rule, the correlation between parents' participation rates and those of the youth is higher for the adolescence periods compared to the two childhood periods. Also, within the adolescence periods, the highest coefficients are those corresponding to late adolescence while they are associated with early childhood, in the case of the childhood periods. Moreover, the coefficients of correlation for youths from two-parent families remain higher than for those emanating from single-parent families. Figure 2 graphs these results.

5 Estimation Results

We now present the estimation results of the bivariate probit model (with repeated observations) discussed in Section 3. Recall that the model allows for the simultaneous estimation of the equation for the participation of young adult and that of his claimant parent. The model also accounts for the unobserved permanent individual characteristics of the parent and the youth, and for their correlation. Finally, a variable (which, in some cases, is vector-valued) of past participation rates of the parents appears as an explanatory variable for the child's participation. The coefficient associated with this variable thus estimates the causal link between parents' and children's participation in the *SA* program.

Since the data are drawn from the administrative records of the *Ministère de la Solidarité Sociale*, family and individual information are only available for periods during which individuals or families receive *SA*. To allow for the absence of continuous data, we set the explanatory variables to be constant across time, and set the parental variables to take the values observed at the end of the claimant parents' last spell on *SA* between 1979 and 1990. After some experimentation, the following variables were included in the vector x_i of exogenous variables (and thus as possible determinants of children's participation):

- The gender of the young adult $(Male = 1)$.
- The number of years of education of the claimant parent $(P - Educ)$. A greater level of education is likely to provide access to better-paid employment for children and may affect their preferences for labor activities [Solon (1992), Gottschalk *et al.* (1994) and

Levine and Zimmerman (1996)]. This variable is also used as a proxy for "potential" income. This is important since, as Levine and Zimmerman (1996) emphasize, the apparent transmission of *SA* receipt may be simply due to an intergenerational transmission of a low earning potential.

- The ethnic origin of the claimant parent $(Canadian = 1)$. The propensity to participate can be expected to vary across ethnic groups [Gottschalk (1996)].
- Primary language ($English = 1$, when the language spoken in the family of origin is English)—this variable may be correlated with possessing the social and economic skills required to integrate into the labor market, as well as with the local characteristics of the environment in which the family lives.
- The region of the claimant parent's residence— average duration on *SA* varies strongly between regions [Duclos *et al.* (1999)]. Thirteen regions were used, with Montreal as the reference region.

All of these variables (except the youth's gender) are also included in the vector X_i of exogenous variables affecting participation in the *SA* system of the claimant parent. The following variables also appear in that vector:

- The number of dependent children $(NChildren)$—Antel (1992) mentions that this variable may affect both the time and money costs of participating in the labor market.
- The age of the claimant parent in January 1990 $(P - Age)$ —individuals aged 46 years and more have longer spells on *SA* [Duclos *et al.* (1999)].
- The household type (single- or two-parent) ($Two - parent = 1$ for two-parent households). A large number of studies have shown that household type has a strong influence on *SA* dependence [Gottschalk *et al.* (1994)].

The exclusion of these three latter variables in the latent equation for the child SA participation insures the identification of the model. Overidentifying tests of these restrictions are reported later in our discussion of the results.

Table 6 presents descriptive statistics on the exogenous variables used in the estimation. In the sample used, 50.3% of children are male. The claimant parents' average level of education is 8.5 years, which is, as expected, not very high. Close to 88% of claimant parents are of Canadian origin.

Table 7 reports the results of two bivariate probit specifications, *e.g.*, a truncated and an untruncated one, for the entire time interval during which the adult children were between 18 to 21 years old. To save space, we only present results for the children's equations. One easily notices that truncating the standard bivariate probit has little impact on the results. Moreover, for both specifications, the presence of unobserved heterogeneity is not rejected. This is revealed by the significant value of the coefficient ρ – which corroborates the hypothesis of a positive correlation between the unobserved characteristics of the parents and of their children that influence participation in the SA program.

Furthermore, the presence of a causal link between parental and youth participation cannot be rejected either. Indeed, the coefficients associated with previous parental participation in Table 7 are statistically significant. They reveal that, at the mean values of the explanatory variables, a one-month increase in parents' participation during the pre-adult phase increases the participation of the youths by 0.1507 month, almost 5 days, during the 48-month observed period (*cf.* last column in Table 7).

Years of parental schooling have, as expected, a negative impact on the youth's propensity to participate. Thus, at the mean of the explanatory variables, an additional year of schooling of the claimant parent reduces the child's participation rate, between the ages of 18 and 21, by 1.4 percentage points. Additionally, the average participation rate of young males is four percentage points lower than that of young females for these same ages. We also notice that being a parent of Canadian origin increases the 18- to 21-year old youth's mean participation rate by eight percentage points. Youths whose parents reside in certain regions, such as Gaspésie, the Saguenay, Mauricie, Estrie and Nord du Québec, have a higher participation rate than those whose parents live in Montreal. Thus, having a claimant parent in Gaspésie at the end of his/her last spell on *SA* increases the participation rate of a 18- to 21-year old youth by 3.4 percentage points over the region of Montreal, at the mean of the explanatory variables.

Table 8 presents the results of a truncated univariate probit model for youths in the age 18–21 interval. This method implicitly imposes that ρ is zero and that there is therefore no correlation in unobserved heterogeneity across generations. Comparing the results in Table 8 with those of the truncated bivariate model for the same ages (Table 7), we see that in both cases the coefficients associated with previous parental participation are positive and significant. However, its value is higher (by 32%) when unobserved heterogeneity is ignored (0.6614 *vs.* 0.5099). Also, the equality of these two coefficients is rejected at the 5% level. This result is explained by the fact that some part of the correlation is now attributed to a causal link while it is actually due to correlation in unobserved heterogeneity. This illustrates the importance of accounting for intergenerational correlation in unobserved individual and environment-specific characteristics.

We can use the results of the bivariate model for the 18–21 interval (Table 7) to illustrate the impact of parental participation on the child's participation over his entire adulthood (18–65). To do this, assume for simplicity that the marginal participation effect remains constant over the life cycle. In this case, the model predicts that a one-month increase in the parent's participation while the youth was aged between 7 and 17 years will increase the child's participation by 1.81 month $(= 0.1507 \cdot (48/4))$ over the ages 18 to 65. This result suggests that a policy intervention targeted at reducing the participation of parents of a 7–17 year-old child will have a non-negligible long-term impact on the child's participation over his entire life cycle.

As discussed above, an interesting issue is whether prior parental participation has different impacts according to the stage of the child's development at which it occurs. To investigate this, total parental participation during the child's pre-adult phase has been broken down into four parts: early childhood (7–9 years), late childhood (10–12 years), early adolescence (13–15 years) and late adolescence (16–17 years). Notice that these four periods are not of the same length (3 years *vs.* 2 years). Consequently, in order to interpret correctly the corresponding coefficients, we transform them so they can be expressed in terms of the total number of months of parental participation in each spell.

The results of these estimates are presented in Table 9. Once again, controlling for selection bias has a negligible impact on the results. Moreover, the four parental participation coefficients have the expected positive sign. We observe that the impact is stronger in early childhood than in late childhood. The impact falls during early adolescence and peaks during the last two years of adolescence. A one-month increase in parental participation induces a 0.2282 month increase in the child's participation between the ages of 18 and 21 if the increase occurs during the 7–9 interval, 0.0242 month for an increase over the 10–12 interval, 0.1587 month for an increase over the 13–15 interval, and 0.2409 month if the increase occurs in late adolescence. A likelihood ratio rejects the hypothesis that the effects of previous parental participation during these four stages are equal (LR test statistic $=$ 536.2767 $>$ 7.8147 $= \chi^2_{(3,\alpha=5\%)}$).

Several explanations for these differences may be advanced. First, as mentioned above, it is possible that, as the child approaches the age of 18 (the age threshold for eligibility to *SA*[7]), the claimant parent of the Quebec SA program encourages the youth to participate in the program in order to partly offset the reduction of the parent's own benefits. Second, the learning effect pertaining to using the program may be more pronounced during the late adolescence stage. Finally, the negative impact of parental *SA* participation on the family stigma felt by a youth (the imitation effect) may be greater when he is still a child (7–9) than during other stages of his development. Nonetheless, according to our results, the causal link between parents' and children's participation is strongest when prior parental participation occurs when the youth is in late adolescence.

The exclusion of the $Nchildren$, $Two - parent$ and $P - Age$ variables in the child's equation ensures that the model is identified. To test for over-identification, we estimated

[7]This threshold applies to all youths except for young mothers, whose threshold is 15 years.

various specifications of the model in which all of these variables were successively paired and added into the equation for the youth's participation in order to leave a unique identifying variable in the parent's participation equation. Table 10 shows test statistics from Lagrange multiplier and likelihood ratio tests. At a 5% significance level, the critical value for the test is 5.9914 ($\chi^2_{(2,\alpha=5\%)} = 5.9914$). All tests fail to reject the over-identifying restrictions for each specification.[8] Moreover, as shown in table 10, adding these variables into the child's equation has little impact on the marginal effects of prior parental participation variables.

6 Conclusion

This paper investigated the intergenerational transmission of reliance on social assistance (*SA*) participation in in Canada. It found evidence of a significant correlation between parental and child participation in *SA*, both during the children's pre-adult and early adult years. This correlation may come from two distinct sources. It may first spring from an imitation and a learning effect. Both establish a causal link between parental and child reliance on *SA*. Or, it may be attributable to the intergenerational transmission of some of the family and environment–specific (observed and unobserved) characteristics that influence *SA* dependence. The policy implications of these two sources of correlation are not the same. When parental *SA* participation *causes SA* participation by children, policies that reduce parents' reliance on *SA* directly reduce future *SA* participation by children. In the absence of a causal link, however, it is not clear whether policies (such as employability programs) that reduce current parental reliance on *SA* would help reduce children's future reliance on *SA*. Instead, severing the intergenerational linkage would then preferably require policies that directly affect the transmission of *SA* determinants.

Our econometric approach allows us to quantify the effect of these two potential sources of intergenerational transmission. Our empirical results confirm the existence of a significant causal link between parental and child reliance on *SA*. We find that, on average, a one-month increase in parental participation during the child's pre-adult period (7–17 years) yields a 0.1507 month (5 days) increase in the participation of the young adult (18–21 years). If we assume that this result obtains over the child's entire period of eligibility, this means that a one-month increment in the parents' participation during the youth's pre-adult stage increases the latter's participation by 1.8 month over the period from 18 to 65. With due account of a proper discount rate, this result implies that any cost-benefit analysis of income security policies should consider not only their impact on the current generation, but also on future generations.

[8] The same procedure was applied to the truncated and untruncated bivariate probit specifications with single parental participation variables and yielded similar results.

We also sought to identify the sub-periods of the youth's pre-adult stage (7–17) during which parental participation had the greatest impact on the child's future behavior. Our results suggest that this impact is greatest during the periods of early childhood (7–9) and late adolescence (16–17). One interpretation of this result is that the imitation effect which structures the youth's preferences is likely to be the strongest when the child is very young while the learning effect dominates when the child approaches adulthood. An additional explanation for the relatively strong effect observed for the late adolescence period is that, in the Québec SA system, claimant parents incur a loss of benefits when the youth reaches 18 years and is not a student. This may induce parents to put some pressure on the youth to claim SA benefits.

Our analysis suffers from a number of limitations which suggest some natural extensions. First, since the data come from administrative records of the Québec *SA* program, all parents used in our analysis have spent at least one month on *SA* during the pre-adult period of the child. Therefore, our results must be properly interpreted as conditional on at least some parental receipt of *SA* (although we do correct for a possible selectivity bias in our estimates). Second, we largely ignore issues of time-dependency in our analysis: more sophisticated multi-spell simultaneous equations transition models would provide an interesting alternative to the repeated-observations bivariate probit model we use in this paper. Third, using samples with youths of various ages would provide an interesting complement to our results based on youths who were 18 years old in 1990. In particular, this would allow us to control for the effect of the overall level of economic activity on intergenerational reliance on *SA*. Finally, the follow-up period could be extended longer into the adult life of the children (*i.e.,* beyond 21 years).

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Table 1
Participation Rates of Four Hypothetical Pairs of Parent-child

	Parental participation rate when their children		Participation rate of the children once they have become adults
	are not yet adults	**have become adults**	
Pair 1	40%	60%	45%
Pair 2	60%	40%	55%
Pair 3	40%	80%	50%
Pair 4	60%	60%	60%

Table 2
Mean Observed Value and Reference Period
for Social Assistance Participation Rates

	Participation Rate	Reference Period
Child Participation variables		
Py18	0.15	12-month period between 1990 and 1991
Py19	0.23	12-month period between 1991 and 1992
Py20	0.27	12-month period between 1992 and 1993
Py21	0.27	12-month period between 1993 and 1994
Py1821	0.23	48-month period between 1990 and 1994
Parent Participation variables		
Pp7	0.37	12-month period between 1979 and 1980
Pp8	0.40	12-month period between 1980 and 1981
Pp9	0.45	12-month period between 1981 and 1982
Pp10	0.53	12-month period between 1982 and 1983
Pp11	0.59	12-month period between 1983 and 1984
Pp12	0.60	12-month period between 1984 and 1985
Pp13	0.60	12-month period between 1985 and 1986
Pp14	0.57	12-month period between 1986 and 1987
Pp15	0.53	12-month period between 1987 and 1988
Pp16	0.50	12-month period between 1988 and 1989
Pp17	0.47	12-month period between 1989 and 1990
Pp18	0.44	12-month period between 1990 and 1991
Pp19	0.43	12-month period between 1991 and 1992
Pp20	0.43	12-month period between 1992 and 1993
Pp21	0.42	12-month period between 1993 and 1994
Pp–	0.51	132-month period between 1979 and 1990
Pp+	0.43	48-month period between 1990 and 1994
Pp79	0.41	36-month period between 1979 and 1982
Pp1012	0.58	36-month period between 1982 and 1985
Pp1315	0.57	36-month period between 1985 and 1988
Pp1617	0.52	24-month period between 1988 and 1990
Pptotal	0.49	180-month period between 1979 and 1994

Table 3
Distribution of Youths by Type of Family of Origin

Type	*Frequency*
Single-parent Family	
Single person	1,267
Single parent (1 child)	4,045
Single parent (2 children)	5,169
sub-total	**10,481**
Two-parent family	
Childless couple	322
Two parents (1 child)	1,564
Two parents (2 children)	4,836
sub-total	**6,722**
Total	**17,203**

Table 4
Coefficients of Correlation between the Annual Participation Rates of Children and those of Parents at the Pre-adult (7-17 years) and Adult (18-21 years) Periods of Children

Participation Rate of Parents *	*Participation Rate of Children*			
	Py18	*Py19*	*Py20*	*Py21*
Pp-				
Total sample (1)	0.2085**	0.2077	0.1992	0.1924
	[0.0001]***	[0.0001]	[0.0001]	[0.0001]
Two-parent families (2)	0.2580	0.2565	0.2492	0.2328
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families (3)	0.1808	0.1761	0.1658	0.1634
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Pp+				
Total sample	0.1978	0.1943	0.1903	0.1846
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Two-parent families	0.2468	0.2421	0.2366	0.2192
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families	0.1680	0.1631	0.1596	0.1600
	[0.0001]	[0.0001]	[0.0001]	[0.0001]

* The family type corresponds to the one observed at the last parental spell on social assistance between 1979 and 1990.
** Correlation coefficient   *** P-value for H_0: correlation coefficient=0
(1) 17,203 obs.   (2) 6 722 obs.   (3) 10481 obs.

Table 5
Coefficients of Correlation between Annual Participation Rates of Children and those of Parents
at Early and Late Childhood (7-9 and 10-12 years)
and Early and Late Adolescence (13-15 and 16-17 years) Periods of Children

Participation Rate of Parents *	*Participation Rate of Children*			
	Py18	*Py19*	*Py20*	*Py21*
Pp79				
Total sample (1)	0.1583**	0.1638	0.1532	0.1450
	[0.0001]***	[0.0001]	[0.0001]	[0.0001]
Two-parent families (2)	0.2016	0.2092	0.1975	0.1876
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families (3)	0.1340	0.1351	0.1243	0.1157
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Pp1012				
Total sample	0.1530	0.1504	0.1450	0.1373
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Two-parent families	0.1987	0.1934	0.1854	0.1664
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families	0.1248	0.1206	0.1163	0.1141
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Pp1315				
Total sample	0.1760	0.1761	0.1694	0.1676
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Two-parent families	0.2198	0.2216	0.2187	0.2024
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families	0.1485	0.1438	0.1338	0.1401
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Pp1617				
Total sample	0.1802	0.1768	0.1732	0.1724
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
Two-parent families	0.2231	0.2187	0.2187	0.2100
	[0.0001]	[0.0001]	[0.0001]	[0.0001]
One-parent families	0.1541	0.1479	0.1414	0.1443
	[0.0001]	[0.0001]	[0.0001]	[0.0001]

* The family type corresponds to the one observed at the last spell of parents on social assistance between 1979 and 1990.
** Correlation coefficient   *** P-value for H_0: correlation coefficient=0
(1) 17,203 obs.   (2) 6,722 obs.   (3) 10,481 obs.

Table 6
Descriptive Statistics on Exogenous Variables

Variable	Mean	Std. Dev.	Minimum	Maximum
Male	0.5037	0.5000	0.00	1.00
Nchildren	1.8608	1.1816	0.00	10.00
P–Educ	8.5360	2.8141	1.00	20.00
P–Age	43.6943	5.6963	33.00	59.00
English	0.0900	0.2862	0.00	1.00
Two-parent	0.3907	0.4879	0.00	1.00
Canadian	0.8794	0.3256	0.00	1.00
Gaspésie	0.0349	0.1835	0.00	1.00
Bas St-Laurent	0.0398	0.1955	0.00	1.00
Saguenay	0.0490	0.2159	0.00	1.00
Québec	0.1217	0.3269	0.00	1.00
Mauricie	0.0767	0.2661	0.00	1.00
Estrie	0.0463	0.2101	0.00	1.00
Montérégie	0.1420	0.3491	0.00	1.00
Lanaudière	0.1141	0.3179	0.00	1.00
Laurentides	0.0964	0.2952	0.00	1.00
Outaouais	0.0503	0.2185	0.00	1.00
Abitibi	0.0312	0.1737	0.00	1.00
Nord	0.0292	0.1685	0.00	1.00

Table 7
Estimation Results for the Child's Equation
Bivariate Probit Models

	Bivariate Probit			Truncated Bivariate Probit		
	Dependent variables: (Py1821,Pp1821)			Dependent variables: (Py1821,Pp1821)		
Variable	Coefficient	Standard error	Marginal effect	Coefficient	Standard error	Marginal effect
Constant	-0.7874	0.0055		-0.7877	0.0056	
Pp717	0.5100	0.0032	0.1508*	0.5099	0.0036	0.1507*
English	-0.0811	0.0038	-0.0247	-0.0810	0.0023	-0.0247
Canadian	0.3083	0.0036	0.0799	0.3083	0.0035	0.0799
Male	-0.1283	0.0028	-0.0398	-0.1282	0.0017	-0.0397
P-Educ	-0.0486	0.0005	-0.0144	-0.0486	0.0005	-0.0144
Gaspésie	0.1205	0.0055	0.0339	0.1207	0.0012	0.0340
Bas St-Laurent	-0.0705	0.0061	-0.0214	-0.0703	0.0017	-0.0213
Saguenay	0.1170	0.0059	0.0330	0.1171	0.0012	0.0330
Québec	-0.1081	0.0041	-0.0333	-0.1078	0.0016	-0.0332
Mauricie	0.0588	0.0049	0.0170	0.0590	0.0015	0.0170
Estrie	0.0452	0.0037	0.0131	0.0454	0.0012	0.0132
Montérégie	-0.0534	0.0033	-0.0161	-0.0532	0.0012	-0.0160
Lanaudière	-0.1137	0.0038	-0.0351	-0.1136	0.0019	-0.0350
Laurentides	-0.1111	0.0043	-0.0342	-0.1107	0.0014	-0.0341
Outaouais	-0.1157	0.0042	-0.0357	-0.1156	0.0011	-0.0357
Abitibi	-0.0170	0.0064	-0.0051	-0.0167	0.0011	-0.0050
Nord	0.1745	0.0038	0.0480	0.1751	0.0022	0.0482
ρ	0.1399	0.0024		0.1399	0.0018	
Log-likelihood	-955799.3140			-955799.2882		

* This marginal effect measures the impact of a one-month increase in parents' prior participation on the number of months of participation by the child when aged 18 to 21, at the mean values of the explanatory variables.

Table 8
Estimation Results for the Child's Equation
Truncated Univariate Probit Model

Variable	Coefficient	Standard error	Marginal effect
Constant	-0.8825	0.0081	
Pp717	0.6614	0.0048	0.1953
English	-0.0829	0.0061	-0.0253
Canadian	0.2797	0.0059	0.0734
Male	-0.1291	0.0031	-0.0400
P-Educ	-0.0450	0.0006	-0.0133
Gaspésie	0.1368	0.0089	0.0382
Bas St-Laurent	-0.0615	0.0088	-0.0186
Saguenay	0.1378	0.0079	0.0385
Québec	-0.0926	0.0061	-0.0283
Mauricie	0.0684	0.0068	0.0197
Estrie	0.0602	0.0082	0.0174
Montérégie	-0.0400	0.0058	-0.0120
Lanaudière	-0.1046	0.0062	-0.0321
Laurentides	-0.0954	0.0065	-0.0292
Outaouais	-0.1024	0.0081	-0.0314
Abitibi	0.0078	0.0097	0.0023
Nord	0.1984	0.0096	0.0540
Log-likelihood	-424812.6391		

Table 9
Estimation Results for the Child's Equation
Bivariate Probit Models with Parents' Participation
broken down into Four Sub-Periods

	Bivariate Probit			Truncated Bivariate Probit		
	Dependent variables: (Py1821,Pp1821)			Dependent variables: (Py1821,Pp1821)		
Variable	Coefficient	Standard error	Marginal effect	Coefficient	Standard error	Marginal effect
Constant	-0.7953	0.0061		-0.7953	0.0026	
Pp79*	0.7729	0.0104	0.2282**	0.7733	0.0099	0.2282**
Pp1012*	0.0821	0.0114	0.0242**	0.0821	0.0080	0.0242**
Pp1315*	0.5375	0.0102	0.1587**	0.5375	0.0087	0.1587**
Pp1617*	0.8167	0.0181	0.2411**	0.8162	0.0134	0.2409**
English	-0.0824	0.0035	-0.0251	-0.0824	0.0015	-0.0251
Canadian	0.3101	0.0041	0.0801	0.3100	0.0031	0.0801
Male	-0.1275	0.0028	-0.0395	-0.1275	0.0028	-0.0394
P-Educ	-0.0476	0.0005	-0.0141	-0.0476	0.0005	-0.0140
Gaspésie	0.1201	0.0050	0.0338	0.1203	0.0029	0.0338
Bas St-Laurent	-0.0709	0.0049	-0.0215	-0.0709	0.0016	-0.0215
Saguenay	0.1168	0.0055	0.0329	0.1168	0.0012	0.0329
Québec	-0.1056	0.0047	-0.0324	-0.1057	0.0016	-0.0325
Mauricie	0.0623	0.0051	0.0179	0.0622	0.0015	0.0179
Estrie	0.0509	0.0044	0.0147	0.0509	0.0013	0.0147
Montérégie	-0.0505	0.0035	-0.0152	-0.0505	0.0012	-0.0152
Lanaudière	-0.1109	0.0036	-0.0341	-0.1106	0.0028	-0.0340
Laurentides	-0.1061	0.0047	-0.0326	-0.1062	0.0013	-0.0326
Outaouais	-0.1137	0.0056	-0.0350	-0.1139	0.0020	-0.0351
Abitibi	-0.0155	0.0057	-0.0046	-0.0152	0.0011	-0.0045
Nord	0.1817	0.0042	0.0498	0.1821	0.0025	0.0499
ρ	0.1257	0.0025		0.1258	0.0026	
Log-likelihood	-955531.1752			-955531.1499		

* Parental participation is expressed in months.

** This marginal effect measures the impact of a one-month increase in parents' prior participation on the number of months of participation by the child when aged 18 to 21, at the mean values of the explanatory variables.

Table 10
Overidentification tests and marginal effects according to various specifications

Variables	Restricted Model	Identifying variable: Two-parent	Nchildren	P-Age
Tests				
Lagrange multiplier		2.4460	0.5706	2.4927
Log-likelihood ratio		2.2528	1.8330	1.9108
Marginal effects				
Pp79	0.2282	0.2193	0.2196	0.2310
Pp1012	0.0242	0.0246	0.0242	0.2310
Pp1315	0.1587	0.1573	0.1624	0.1584
Pp1617	0.2409	0.2464	0.2788	0.2123

Figure 1: Correlations between Annual Participation Rates of Children and Anterior and Future Participation Rates of Parents

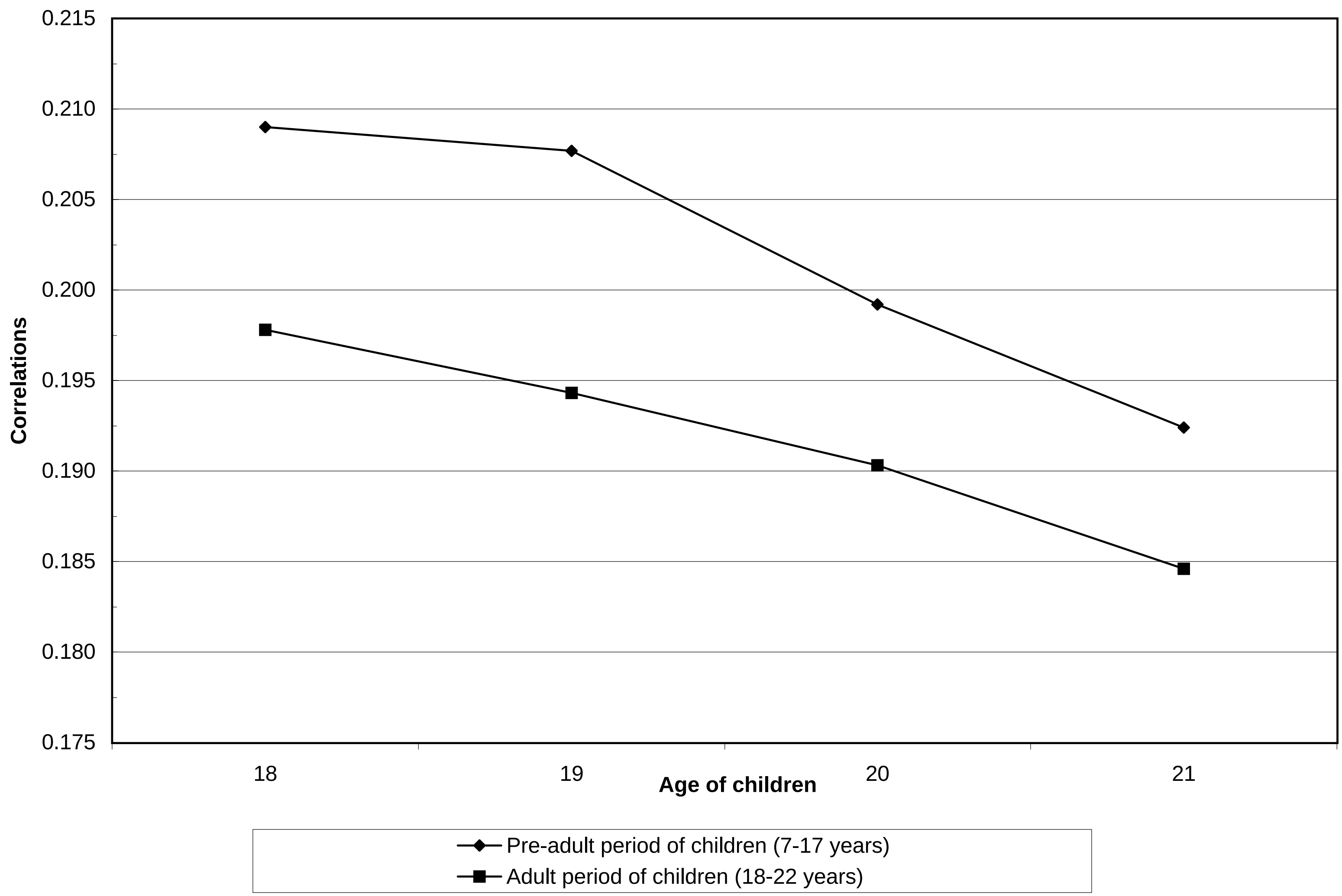

Figure 2: Correlation Coefficients between Annual Participation Rates of Children and that of Parents during Pre-adolescence and Adolescence Periods of Children