DISCUSSION PAPER SERIES

No. 7299

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PUBLIC POLICY

Centre for Economic Policy Research

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Available online at: www.cepr.org/pubs/dps/DP7299.asp

ISSN 0265-8003

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Discussion Paper No. 7299
May 2009

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CEPR Discussion Paper No. 7299

May 2009

ABSTRACT

Public School Availability for Two-year Olds and Mothers' Labour Supply

French children start public school either the year they turn two or the year they turn three. We evaluate the impact of this unique schooling policy on maternal labour supply. Using a Regression-discontinuity design, we show that early school availability has a significant employment effect on lone mothers, but no effect on two-parent families. Also we show that the effect grows larger as the child grows older and as the family loses eligibility for child benefits. Finally, we provide some new evidence that school enrolment at the age of two has no adverse effect on children's subsequent educational outcomes.

JEL Classification: J13 and J22
Keywords: maternal labor supply and preschool

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Submitted 05 May 2009

Given many recent policy proposals to extend pre-school programs across developed countries, it is important to understand what the consequences of these policies will be. The first expected outcome of such policies is to enhance child development and improve their subsequent adaptation to elementary school. Another expected outcome is to provide families with free child care and to make it easier for parents with young children to participate in the labour market.

Several recent studies have explored the effect of preschool availability for five-year olds or even four-year olds[2], but little is known about what the effect of expanding eligibility to three-year or two-year olds would be. In particular, it is not known whether it would really contribute to increase the participation in the labour market of mothers, especially lone mothers, and help reducing poverty.

Within this context, France represents a very interesting case since it introduced universal pre-elementary school for three-year olds in the mid nineties. About one third of French children start public school even earlier, at the age of two[3]. French pre-elementary schools are genuine part of the primary school system. The curriculum is defined at the national level and teachers are certified primary school teachers. The program consists of 28 hours per week during the same academic year as elementary schools. The annual cost per pupil is about 4,700 euros. This system provides a unique opportunity to study the causal effects of public school availability on families with children aged two or three.

To identify the effect on labour market participation, we make use of the highly discontinuous relationships between children's exact date of birth and pre-elementary school eligibility. French academic year begins in September and children can start school either in September

[2] See for example, Baker, Gruber and Miligan (2008), Cascio (2009a), Fitzpatrick (2008a), Berlinski and Galinani (2007), Schlosser (2005).

[3] According to European Commission (1995), France and Belgium are the only countries where children can start school as early.

of the year they turn three (normal start) or September of the year they turn two (early start, one third of a birth cohort). This regulation generates very significant discontinuities in early school enrolment between children born in December and children born at the beginning of January of the following year, both within the group of two-year olds and within the group of three-year olds. For example, according to the census conducted in 1999, almost all children born in late December 1996 started school at the beginning of academic year 1998/1999, against only about 75% of those born in early January 1997.

Given this fact, the first basic question is whether we observe similar discontinuities in the relationship between mothers' labour market participation and their children's exact date of birth, which would suggest a causal relationship between access to preschool and mother's labour force participation. Interestingly, we find no discontinuities for two-parent families, but significant ones for single-mothers. According to the 1999 census, there exists a four percentage point variation in participation rate between lone mothers whose youngest child was born just before the 1st January of 1997 and lone mothers whose youngest child was born just after this date. Further investigations reveal that this effect varies a lot across sub-populations of single mothers. In particular, the estimated effect of preschool availability is twice as large when the child is three years old as when she is only two years old. This finding is consistent with the fact that non-working mothers lose their eligibility to family benefits when their child turns three. French mothers of two-year olds are still eligible to a benefit of about half the minimum wage and our analysis reveals that it divides by two the effect of free pre-elementary school availability on their propensity to work.

Overall, public pre-elementary school represents a mode of child-care which is not more costly for the government than parental care[4], but which increases significantly the ability of lone mothers to work. Given these results, further expansion of pre-elementary schools for

[4] The public aids given to parents who stay at home with children aged 3 or less represent about 5,000 euros per child and per year, which is very close to the cost of pre-elementary schools.

two-year olds at the detriment of parental care would be unambiguously cost-effective provided that it has no adverse effects on children's subsequent outcomes. The question of whether children can start school as early as at the age of two is highly controversial and it is beyond the scope of this paper to address this issue in detail. Building on the specificities of the French pre-elementary admission rules, we provide nevertheless some new simple evidence suggesting that, if any, the effect of substituting pre-elementary school for parental care on two-year olds' subsequent outcomes is weak. In particular, we show that official regulations generate very significant variation across month of birth in the difference in early school enrolment across French regions, but no variation at all of the difference in children's subsequent educational outcomes.

The paper is organized as follows. The first section provides an overlook of the related literature. The institutional context and the data used are presented in the second section. We use the full census records which makes it possible to identify variations in the relationship between children's exact date of birth and families' outcomes just before and after the cut-off dates for early entry into school. The empirical analysis of the effect of early enrolment is developed in section III and section IV. The last section provides additional evidence on the effect of early enrolment on educational outcomes.

I. Related Literature

Several recent studies exploit variations in the presence of public preschools across geographic areas over time in order to identify the effect of these programs on maternal employment. For example, Cascio (2009a) argues that the introduction of kindergartens for children aged five in the US public schools during the 1960s and 1970s raised very significantly the labour supply of single mothers with eligible children, but had no effect on other mothers. In a related paper, Baker et al. (2008) show that the extension of full-time kindergartens to all five year olds (and the provision of childcare at a price of 5 dollars per day to all children aged 4 or less) in the Canadian province of Quebec in the late nineties coincides with an increase in maternal labor supply and a decline in children's outcomes in this province (compared with the rest of Canada). One issue with this evaluation is that there were several other changes to the benefits paid to families, both in the province of Quebec and the rest of Canada in the late 1990s (as well as other specific reforms in other provinces).

Berlinski and Galiani (2007) analyze the impact of a program of construction of pre-elementary schools in Argentina. Using the difference in the timing of construction across regions, they find a significant effect of the program on both preschool enrolment en maternal employment. According to Berlinski, Galiani and Gertler (2009), this program also had a positive effect on pupils' outcomes. Schlosser (2005) evaluates the impact of a reform conducted in 1999 in Israel which increased the availability of free preschool for children aged 3 and 4 in the poorest towns of the country. Building on the difference in the timing of the reform across municipalities, she finds that free preschool availability increased both preschool attendance and maternal labour supply in Arab municipalities.

In a very different contribution, Gelbach (2002) shows that the quarter of birth of five year olds affects maternal employment and children's preschool attendance in a parallel way. Using quarter of birth as an instrument, his analysis suggests that preschool availability has a significant effect on maternal employment. It is not clear, however, whether the effect of quarter of birth on mothers' behaviour is due to free child care availability only. Children born later in the year are also younger at each point of time, which may, as such, be an explanation for the lower participation of their mother in the labour market. Finally, the recent paper of Fitzpatrick (2008a) examines how universal pre-kindergarten availability for 4-year olds affects mothers' labour supply in Georgia and Oklahoma. Comparing children born just before and just after the eligibility cut-off (i.e., September 1st), she does not find any robust impact on mothers' behaviour, except some minor changes in rural areas.

To the best of our knowledge, our paper is the first to address whether the availability of formal classroom-based learning environment for 2-year and 3-year olds affect families' outcomes.

II. French Institutional Context and Data

Before starting school, the vast majority of French families have either to stay at home with their children or to use informal child care arrangements (see e.g. Blanpain, 2006). As it happens, the enrolment capacity of alternative publicly-funded child care arrangements is limited[5]. Also these options are far from fully subsidized and the private costs for families is non negligible[6]. Within this context, any increase in free public school availability reduces very significantly the costs of child-care and increases the net potential wages of mothers. Our basic research question is whether it affects maternal labour supply. Our empirical strategy makes use of the specific admission rules of the French pre-elementary school system and this section begins by describing these rules.

II.1 . Eligibility rules

French academic year begins in September and children start pre-elementary school either in September of the year they turn three (normal start) or in September of the year they turn two (early start). Early starters (about one third of a birth cohort in the late nineties) spend four years in pre-elementary school from year 1 (called *Très Petite Section*) to year 4 (called

[5] Only about 10% of children can be enrolled in collective child care centres (called *crèche*) and only about 20% can benefit from a registered child-minder (called *assistante maternelle agréée*).

[6] The average costs for families is about 300 euros per month, i.e., about one third of the minimum wage.

Grande Section). The other children spend only three years in pre-elementary school from year 2 (*Petite Section*) to year 4. All children start elementary school in September of the year of their sixth birthday[7].

Within this system, the school enrolment rate observed in March of any calendar year *t* is close to zero for children born in *t-2*. They were too young by the start of the current academic year (September of year *t-1*), even for an early enrolment. In contrast, the enrolment rate is close to 1 for children born in *t-4*. They have all benefited either from a normal start by the beginning of the current academic year or from an early start by the start of the previous academic year. Lastly, the enrolment rate lies somewhere between 0 and 1 for children born in *t-3*. A fraction only has benefited from an early start at the beginning of the current academic year. Overall, at the beginning of any calendar year, eligibility rules generate two potential discontinuities in school enrolment rates, one between children born in late *t-4* and children born in early *t-3* (the three-year olds 'group) and one between children born in late *t-3* and children born in early *t-2* (the two-year olds' group). The basic questions asked in this paper will be whether these discontinuities are actually perceptible in our dataset and whether they coincide with discontinuities in the variation of maternal participation in the labour market. Another important issue will be whether the magnitude of the discontinuity is the same for the two-year olds' group (child born in *t-4/t-3*) as for the three-year olds' (*t-3/t-2*). In 1999, mothers who do not work are actually eligible to a benefit of about 530 euros (more than half the minimum wage) only until their youngest child turns three and the question is whether it affects their response to public school availability.

As it happens, the participation decision of two-year-old' s mother is not simply a matter of comparing her net hourly wage with the value of an additional hour of non-paid activities

[7] In France, pre-elementary schools and elementary schools belong to the same public system of *primary* education. The curricula of both pre-elementary and elementary schools are defined at the national level by the same administration. Teachers ("professeurs des écoles") are the same in both types of schools. Any given *professeur des écoles* can teach either in pre-elementary or in elementary schools and move from one type of schools to the other.

(“leisure”), it is also a matter of whether the family benefits alone do not provide her family with a better standard of living than the one that can be reached through labour market participation and paid-work. If family benefits are relatively high compared to potential labour earnings, the availability of public school may have no effect on a mother’s labour supply even when it increases net hourly wage above the value of an additional hour of non-paid activity. Under this assumption, the effect of public school availability on mothers’ participation in the labour market is likely to be more significant when children are three years old than when they are only two years old and their mothers still eligible to benefits.

II.2 Data

This paper uses the population census conducted in March 1999. Specifically, we use the full census database. It provides us with exhaustive information on the labour market and family situation of the total population of French mothers as well as on the date of birth, number and sex of their children. In March 1999, children born in 1997 are still too young to have benefited from an early enrolment at school, whereas a fraction only of those born in 1996 have benefited from an early start. To analyse the enrolment and participation discontinuities generated by French preschool regulations, we will mostly focus on the population of mothers whose youngest child was born either in December 1995-January 1996 or in December 1996-January 1997. Each of these specific segments of the population corresponds to about 90,000 families (10,000 lone-parent and 80,000 two-parent families). It is only because we have access to the full census records that we can focus on such narrow bandwidths and detect the variation in families’ behaviour just before and after the cut-off dates for early entry into public school.

III. Graphical Analysis

This section provides some graphical evidence on the relationships between exact date of birth, preschool attendance and maternal participation in the labour market, as measured in March 1999 in the general census of the population. We focus on mothers whose youngest child was born during either the month before or the month after the cut-off entry date (i.e., 1996, January, 1st). Also, we provide a separate analysis for two-year olds and three-year olds as well as for mothers in two-parent families and single-parent families. Single-parent families are faced with financial and time constraints that are much more severe than those of two-parent families and are likely to be much more directly affected by free public school availability.

To begin with, Figures 1a and 1b show the relationship between exact date of birth and pre-elementary school attendance in March 1999 for mothers whose youngest children were born between December 1995 and January 1996. These children are three years old by the time of the census. Figure 1a focuses on mothers in two-parent families and Figure 1b on single-parent families. Comfortingly, the graphs reveal a very significant drop in pre-elementary school attendance rate between late December 1995 and early January 1996. This pattern is consistent with the eligibility rules where pupils can start either in September of the year of their third birthday (normal start) or in September of the year of their second birthday (early start). The vast majority of children born in late 1995 have benefited from either a normal start (in September 1998) or an early start (in September of 1997) whereas a fraction only of those born in early 1996 have benefited from an early start (in September 1998).

Figures 2a and 2b focus on the same samples of mothers and show the relationship between children's exact dates of birth and mothers' labour market participation. Interestingly, Figure 2a does not show any discontinuity between December 1995 and January 1996 for two-parent families. In contrast, Figure 2b reveals a clear discontinuity for single mothers (about 4 points). As discussed above, pre-elementary school availability increases potential net wages, but the effect on labor market participation is perceptible for lone mothers only.

Figures 3a and 3b show the relationships between exact date of birth and pre-elementary school attendance in March 1999 for children born between December 1996 and January 1997. These children are two years old by the time of the census. Unsurprisingly, the graphs show a drop in pre-elementary school attendance between late December 1996 and early January 1997 which is almost as significant as the drop observed in Figures 1a and 1b. The attendance rate is close zero for children born in early 1997 whereas it is above 15% for children born in late 1996. Children born in early 1997 were too young in September 1998, even for an early start, whereas a significant fraction of those born in 1998 have actually benefited from an early start.

Figures 4a-4b focus on the same samples of mothers as Figures 3a-3b and show the relationships between children's exact dates of birth and mothers' labour market participation. Figure 4a does not show any discontinuity for two-parent families, whereas Figure 4b reveals a discontinuity for single mothers (about 2 points). Hence, the pattern is similar to Figures 2a-2b, even though the single-mother discontinuity is less clear-cut. As discussed above, the smaller effect for mothers of two-year olds is consistent with the fact that those who do not work are still eligible to family benefits.

Overall, the graphs suggest that the same discontinuities exist in the relationships between single-mothers labor market participation and their children's dates of birth as in the relationships between early school enrolment and children's dates of birth. In the next section we discuss the conditions under which this may be interpreted as reflecting a causal effect of early school availability on labor market participation and develop the corresponding regression analysis.

IV. Regression analysis

For child i, we denote D_i the exact date of birth and S_i a dummy indicating actual school enrolment. Also, following Imbens and Angrist (1994) and Hahn, Todd and Van der Klaauw (2001), we denote $S_i(d)$ the potential school enrolment status of child i given any date of birth d. We assume that $S_i(d)$ is independent of D_i and non-increasing in d for each i. Lastly we define L_{i1} the potential labour market participation of i's mother if i has already started school and L_{i0} the same potential outcome if i has not yet started school.

We want to evaluate the distribution of the causal effect L_{i1}-L_{i0} of school enrolment on mothers' participation in labour market and the fundamental issue is that we never observe simultaneously L_{i1} and L_{i0}. Our identifying strategy exploits the discontinuities in the probability of enrolment observed at d_c=January 1st 1996 or at d_c=January 1st 1997.

To be more specific, let us denote ΔL_c the discontinuity in actual participation rate at d_c,

$$\Delta L_c=\lim_{d\downarrow dc}[\mathrm{E}(\mathrm{L}_i/D_i=d)-\mathrm{E}(\mathrm{L}_i/D_i= d_c)],$$

and ΔS_c the discontinuity in school enrolment rate at the same cut-off date,

$$\Delta S_c=\lim_{d\downarrow dc}[\mathrm{E}(S_i/D_i=d)-\mathrm{E}(S_i/D_i= d_c)].$$

Under the assumptions that the potential outcome $\mathrm{E}(L_{i0}/D_i=d)$ is continuous at d_c and the causal effect L_{i1}-L_{i0} is independent from date of birth D_i, the ratio between these two discontinuities identifies the average causal effect of school enrolment on families whose enrolment status change at d_c (see Hahn, Todd and Van der Klaauw, 2001),

$$\Delta L_c / \Delta S_c = \lim_{d\downarrow dc}[\mathrm{E}(L_{i1}-L_{i0}/S_i(d)-S_i(d_c)=1;D_i =d)].$$

In substance, our identifying assumptions mean that the participation rates would be continuous at d_c if enrolment status did not change at d_c so that any discontinuous shift in actual participation rates isolate the causal effect of change in school enrolment status. Note that these assumptions would plausibly be violated if mothers were able to manipulate exact date of birth in order to modify the school entry age of their children. Also our assumptions would be problematic if the subsequent fertility of a mother (as well as her probability of becoming a single-parent mother) was affected by her child's exact date of birth (i.e., by whether the birth took place just before or just after the school enrolment cut-off). One simple way to test these assumptions is to test whether there are discontinuities in the distribution of youngest children's date of birth at the cut-off dates. We have checked that it is not the case. Figures A1 and A2 in the Appendix show that there are no discontinuous variations in the number of single-parent families across youngest children's date of birth. We have also checked that when we focus on families with one child born between Dec. 1995 and Jan. 1996 (or Dec. 1996 and Jan. 1997), the probability of being a single-parent family and the probability of having a younger child were both continuous at the cut-off dates (see Figures A3-A4 and A5-A6 in Appendix). These results clearly rule out the assumption that mothers' fertility (or family composition) is affected by whether the exact date of birth is above or below the cut-off date.

With respect to estimation, we need to estimate the ratio of two differences, $\gamma=\Delta L_c/\Delta S_c$. As discussed by Hahn et al. (2001) or Imbens and Lemieux (2008), the simplest way is to use local linear regressions. It amounts selecting the observations within a distance h on either side of d_c and estimating the effect of S_i on L_i by two-stage least squares method with the discontinuity dummy $1[D\geq d_c]$ as the excluded instrument and with a spline function of D (with a knot at d_c) as a control variable. With respect to the bandwidth size h, we will use h=one month as a minimal benchmark and test the robustness of our results to alternative (larger) bandwidth choices.

IV.1 Basic Regression Results

To begin with, let us consider mothers whose youngest child was born between early December 1995 and late January 1996. All these children are three years old in March 1999 (census date) but those born in late 1995 are in the year of their fourth birthday whereas those born in early 1996 are still in the year of their third birthday. As a consequence, all those born in late 1995 have already started school whereas a significant fraction of those born in early 1996 have still not started. Tables 1a and 1b (columns 1 and 2) show the result of regressing children's enrolment at school and mothers' participation in the labour market on a dummy indicating whether the children were born in 1996 (i.e., $1[D \geq d_c]$), where d_c=January 1st, 1996) and a spline function of exact date of birth with a knot at $D= d_c$ (i.e., holding D and the interaction between D and $1[D \geq d_c]$ constant). Table 1a focuses on two-parent families whereas Table 1b focuses on single-parent families. As expected, the first stage analysis confirms that there is a very significant discontinuity in school enrolment probability at d_c. The discontinuity is almost as large for single mothers (14 percentage points) as for mothers in two-parent families (17 percentage points). The reduced form equation does not show any significant discontinuity at d_c in labour market participation for mothers in two-parent families, but reveals a 3.6 percentage point discontinuity for single mothers. When we focus on single mothers, the ratio of the participation and enrolment discontinuities provides an estimation of the average effect of school enrolment at age 3 on maternal participation of about .25 (significant at standard levels)[8]. This Regression-Discontinuity estimate is larger than the corresponding OLS estimate, but the difference is not statistically significant.

[8] Simple interpretation for smaller OLS estimate is that there exist errors in the measurement of school enrolment in the census database which generate an attenuation bias. Another possible interpretation is that Regression-Discontinuity estimates identify the effect of public school availability on families whose school enrolment status is actually affected by public school availability (compliers) whereas OLS estimate provides an estimate of

A similar pattern emerges from the analysis of the sample of mothers whose youngest child was born between early December 1996 and late January 1997 (see Tables 1a and 1b, columns 3 and 4). All these children are two years old in March 1999, but those born in 1996 are in the year of their third birthday whereas those born in 1995 are still in the year of their second birthday. As a consequence, none of those born in early 1997 are at school whereas a significant fraction of those born in late 1996 have benefited from an early start in pre-elementary school.

As expected, the first stage analysis confirms that there is a significant discontinuity in school enrolment probability at d_c =1996 January 1st. It is almost as large as the discontinuity at d_c=1995 January 1st. Also the reduced form analysis confirms that no discontinuity in maternal participation exists at d_c for two-parent families, whereas a 1.4 percentage point discontinuity is perceptible for single mothers. It is not significant at standard levels however. The ratio of the participation and school enrolment discontinuities provides an estimation of the average effect of school enrolment at age 2 on single mothers' participation of about .11 which is not significant at standard levels either. As discussed above, one explanation for this less significant labour supply elasticity at age 2 lies in the fact that mothers who do not work are still eligible to family benefit when their youngest child is still below the age of three.

Overall, Tables 1a and 1b confirm the graphical evidence and suggests that availability of free public school has a significant effect on single mothers, especially after their youngest child turns 3, whereas it has no perceptible effect for mothers in two-parent families, regardless of whether we focus on families with two-year olds or with three-year olds.

As a robustness check, Table A1 in Appendix shows the variation in our regression-discontinuity estimates when use a larger bandwidth (October-March) and control for higher

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the average effect. Higher RD estimates may simply reflect that compliers are characterized by higher labour supply elasticity.

degree polynomials in exact date of birth. Generally speaking, these regressions confirm that the labour supply elasticity is statistically significant for the group of three-year olds (and about 0.35) whereas it is only marginally significant for the group of two-year olds (and about 0.15 only). As an additional check, we have replicated our analysis on the sample of lone mothers whose youngest child was born in November-December 1995 (or November-December 1996) using December 1st as cut-off. These "placebo regressions" yield non-significant regression-discontinuity estimate of the elasticity of mothers' labour supply. The same result holds true when we use the January-February 1996 (or the January-February 1997) sample with the February 1st cut-off.

IV.3 Alternative evidence on maternal labour supply elasticity

French pre-elementary school is free, available to all families (regardless of income). It runs four days a week (plus Saturday morning) for the length of school year and lasts 6 hours per day (typically from 8.30 to 11.30 in the morning and 13.30 to 16.30 in the afternoon). Also, most schools offer extended hours after 16.30 and during the lunch time (lunch being highly subsidized). Overall pre-elementary schools care for children for a large part of the week and represent a very significant subsidy for childcare. Given this reality, it is not clear why the employment effect on two-parent families is negligible.

The population of married mothers who do not work has changed and declined a lot over time and one potential explanation is simply that the labour supply elasticity of this population is now much lower than it used to be in the sixties or the seventies. As it turns out, the recent evaluations of the 1994 reform of family benefits suggest that it is not the case (see e.g. Piketty, 2005). Before 1994, mothers with three children or more received about 550 euros per month (about half the minimum wage) provided that they do not work and that their youngest child is below the age of three. The 1994 reform extended the program to two-child mothers and Piketty (2005) shows that it was followed by a huge drop in the labour market participation of these women. Figure 5 provides new evidence on this issue. Specifically we use the 1999 census and compare the labour market participation of mothers in two-parent families depending on whether their youngest child is just above or just below the 3 years old threshold. The graph reveals a very clear discontinuity of about 4 percent points at the cut-off[9]. A detailed analysis of this effect is below the scope of this paper but it clearly rules out the assumption that the elasticity of female labour supply is negligible in France. Overall, the most plausible explanation for the weak employment effect of preschool on two-parent families is simply that it crowds out alternative non-family childcare. Two-parent families are much less exposed to poverty and have larger social networks than lone mothers[10]. Two-parent families have a much larger access to non-parental modes of childcare (formal and informal) and are likely to use preschool more often as a mere substitute for these non-parental childcare. It is plausibly why their labour supply is much less affected than those of lone mothers.

V. Discussion

As discussed above, the majority of French two-year olds are either at home with their parents or in pre-elementary schools. Given parental benefits, these two modes of child-care represent about the same public costs for the government, i.e., about 5000 euros per child and per year. In other words, the French government invests about the same amount of public money simultaneously into two very different technologies of child-care and one interesting policy question is whether it is cost-efficient, i.e., whether it would not be cost-effective to invest

[9] We have checked that the discontinuity is as large for single-parent families as for two-parent ones.
[10] According to the French National Institute, the poverty rate for single-parent families is about 30% whereas the poverty rate for two-parent families is about 10% (see Goutard and Pujol, 2008).

more in one technology at the detriment of the other one. Given that the two alternative arrangements are free and do not involve private costs for families, the question boils down to compare their social benefits, i.e., their effect on labour market and educational outcomes.

With respect to labour market outcomes, investment in pre-elementary school availability seems clearly a better option since it increases significantly lone mothers' ability to participate in the labour market. Given that single-mothers represent about 12% of the French population of mothers and that the elasticity of their labour supply to pre-elementary school is about 0.25, the value of substituting pre-elementary school for alternative childcare corresponds on average to about +3% of the average annual value of one job per child and per year[11] ($0.03=0.25\times 0.12$).

Given this fact, substituting pre-elementary school for parental child-care would be unambiguously cost effective provided that it has no adverse effect on very young children's subsequent outcomes[12]. Existing statistical evidence on this issue suggests that children who start school at the age of two actually perform better in elementary school than those who start at age three and spend one year less in pre-elementary school (see e.g. Caille, 2001). As usual, this type of result is interesting, but it is not clear whether it really captures the causal effect of starting school by the age of two or a selection bias. One simple way to address this question is to use the fact that the difference in early enrolment rates across regions is much more significant for children born in the middle of the year than for children born either at the beginning or at the end of the year. As it happens, French early enrolment regulations ask schools' heads to give a priority to children born at the beginning of the year (i.e., the most

[11] Assuming that the social value of a job in the French economy is about 40,000 euros per year, the social value of the substituting preschool to parental child-care is about 1,200 euros per child and per year. We should also take into account the discounted effect on mothers' careers. Long disruptions have plausibly a very significant negative effect on subsequent wage mobility (Bayet, 1997).

[12] In the US case, recent research on the effect of prekindergarten on subsequent outcomes include Figlio and Roth (2009), Fitzpatrick (2008b), Magnuson, Ruhm and Waldfogel (2007), Dhuey (2007). Using the timing of school holidays in Dutch primary schools, Leuven, Lindhal and Oosterbeek (2006) provide an analysis of the effect of time in school on 4 years old's outcomes.

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mature of their year group by the start of the academic year) and, consequently, these children have relatively high early enrolment rates in all regions. For symmetrical reasons, children born at the end of the year have relatively low early enrolment rate in all regions. In contrast, children born in the middle of the year have relatively high early enrolment rate in the least constrained regions only, i.e. regions with relatively high overall enrolment capacity only[13]. Figures 6a and 6b show the variation across exact dates of birth in the difference in early enrolment rates across regions with relatively high overall enrolment capacity (i.e., above the median) and regions with relatively low overall enrolment capacity (i.e., below the median). It confirms that the regional differences are actually much more significant for children born in the middle of the year than for children born either at the end or at the beginning of the year. The same strictly convex pattern is observed in the census conducted in 1982 for cohorts born in the late seventies as in the census conducted in 1999 for cohorts born in the late nineties. For example, in 1982, the difference in early enrolment rates across regions is about 30 percentage points for children born in June, whereas it is only about 15 percentage points for children born in late December and about 18% for children born in early January. Given this specific feature of the French enrolment rules, the question becomes whether regional differences in children's subsequent outcomes are themselves more significant for children born in the middle of the year.

The general census of the population makes it possible to address this question using one very basic educational outcome - the probability of early dropout from school. In France, school is compulsory until the age of 16 and the census provides a direct measure of the proportion of individuals who dropout from school at age 17 as well as a measure of the variation in these proportions across regions and exact dates of birth. Do we observe any significant regional

[13] There exists significant variation in public schools' overall enrolment capacity across regions (see e.g. Clément and Nicolas, 2003). In particular, enrolment capacity tends to be lower in more urban areas (such as Parisian region). Our identifying restriction is not that the factors which explain a region's overall enrolment capacity are uncorrelated with the factors which explain performance at school. Our identifying restriction is simply that if this correlation exists, then it does not vary significantly across month of birth within the year.

difference in early dropout rates and is it more significant for individuals born in the middle of the year? Interestingly enough, Figure 7 reveals that the early dropout rates are slightly larger in low enrolment regions, but the difference between high and low enrolment regions is about exactly the same for students in the middle of the year as students born either before or after this period. In fact, in both regions, the early dropout rates declines in a highly parallel way from about 9 percentage points for pupils born in January to about 5 percentage points for pupils born in December, without any clear fluctuations in the regional difference across days of birth within the year[14].

It is possible to put this graphical analysis into a regression format and it confirms that there is no statistically significant variation in the effect of day of birth within the year across regions, which is consistent with early start having no significant negative effect on subsequent outcomes[15].

We have also checked that the same result holds true when we focus on the scores obtained in primary school at the national test conducted at the beginning of the third grade. The difference in test scores across high and low enrolment region is small (about 6% of a SD) and it is almost exactly the same for pupils born in the middle of the year or either at the beginning or the end of the year (available on request).

Early start policy is often presented in the French debate as a relatively costly policy option. Our analysis suggests that it is not the case. It increases maternal ability to participate in the

[14] The same result holds true when we use the proportion of children who dropout from school at age 18 or at age 19 as educational outcome (rather than the proportion who dropout from school at age 17).

[15] This analysis is available on request. Using jointly 1982 and 1999 censuses, it is actually possible to construct a two-sample instrumental variable estimate of early start. First we perform the first-stage analysis with the 1982 census, i.e. the regression of early start on region, date of birth and the interaction between region and date of birth. Secondly, we use the results of this initial regression to construct predicted early enrolment rates with the 1999 data and we perform the second-stage analysis with these predicted values (using the interaction between being born in the middle of the year and living in a high-enrolment region as excluded instrument). Standard deviation has to be adjusted, see Angrist and Krueger, 1992. Under the maintained assumption that the effect of early start does not vary significantly across regions, this approach provides an estimate of the causal effect of early start on early dropout which turns out to be small and not statistically significant.

labor market without having adverse effect on children and without involving more public spending than the other aided options[16].

VII. Concluding remarks

French children start pre-elementary school either the year they turn two or the year they turn three. This paper provides the first comprehensive evaluation of the effect of this unique early schooling policy on maternal labour supply. We find that pre-elementary school availability has significant employment effect on single-parent families, but no effect on two-parent families. Two-parent families have a larger access to alternative non-parental modes of childcare and it is likely that pre-elementary school crowds out these non-parental alternatives. It is also highly plausible that the substitution of free pre-elementary school for alternative (non-free) modes of childcare increases family net resources and well-being, even though it is not perceptible in our data. Further research is needed in this area to better understand the benefits of free pre-elementary schools for families.

More research is also needed to better explore the effect on children of starting school as early as at the age of two. French early start regulations generates very significant variations across month of birth in the difference in early start probability across regions and we find that these variations do not coincide with variation in subsequent educational outcomes. It suggests that starting school at the age of two has no adverse effect on subsequent outcomes (which is consistent with existing statistical analysis), but it would be obviously very important to have additional evidence relying on alternative source of identification.

Taken together, our results suggest that early school policy is more cost-effective than existing alternative subsidized modes of childcare for two-year olds in France. It does not

[16] It does not mean that it is necessarily more cost effective than any other comparison policy. If the only goal is to increase maternal employment, it would be plausibly even more cost effective to eliminate the current subsidy to mothers not working (or to combine this elimination with the development of pre-elementary schools).

mean that it is the most cost-effective institution one can possibly conceive. Further research is needed to explore this issue and help designing better preschool institutions.

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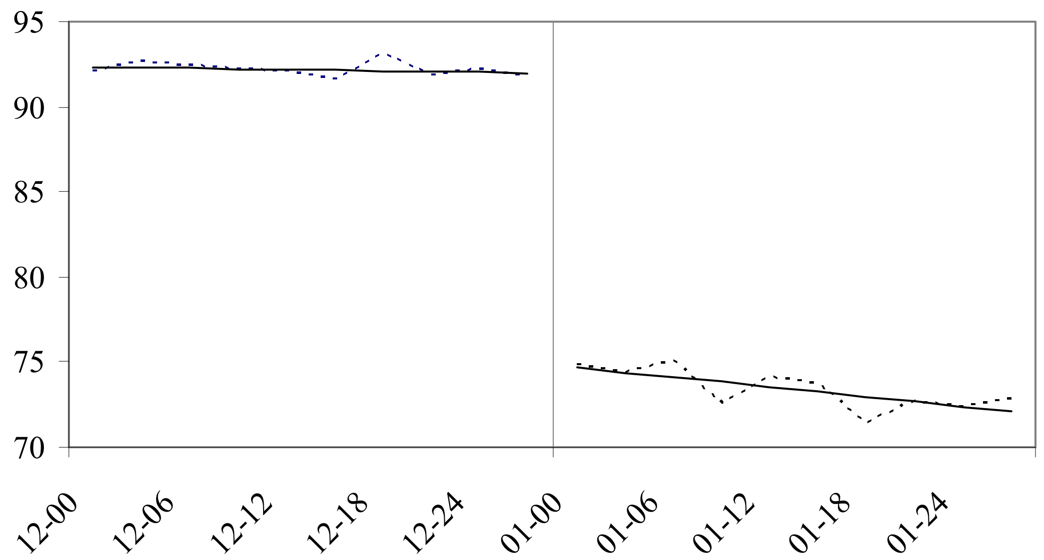
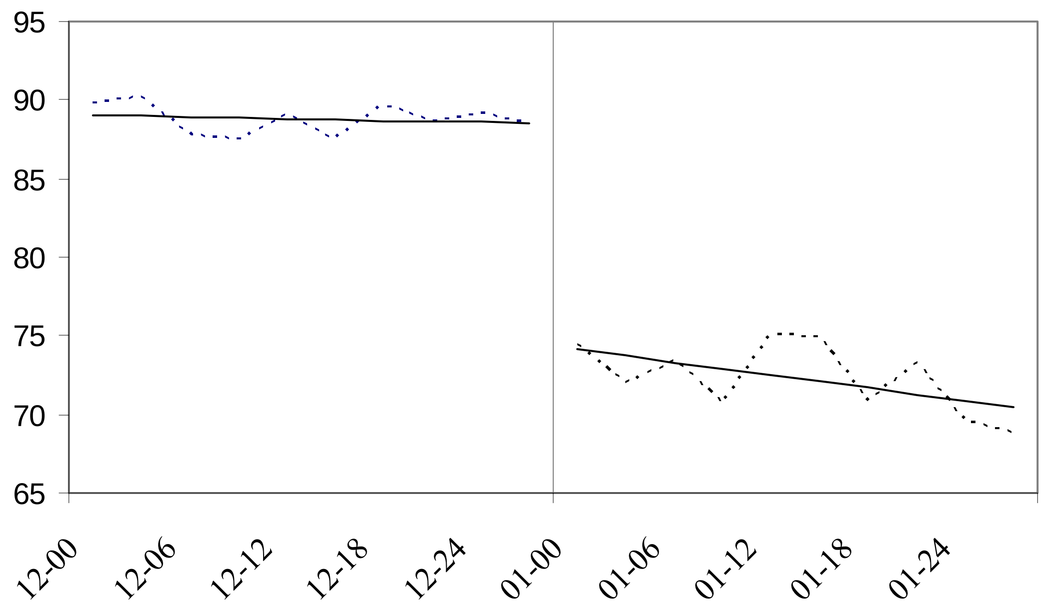
Figure 1a: School enrollment, two-parent families (1995-96 cut-off, age 3)



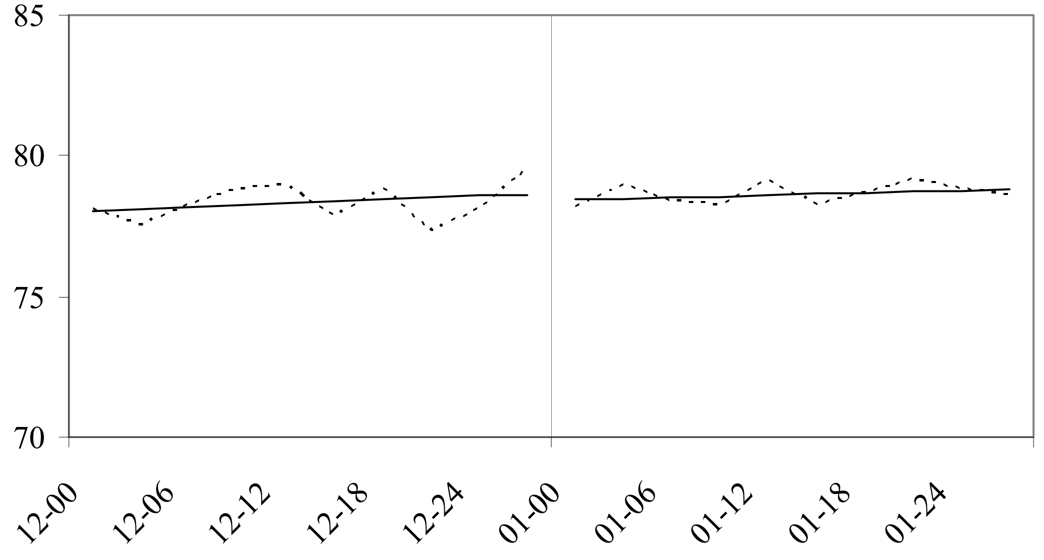
Source: 1999 census.
Sample: Two-parent families whose youngest child was born between Dec. 1995 and Jan. 1996.
Reading: The enrolment rate is above 90% for children born in Dec. 1995, it is below 75% for children born in Jan. 1996

Figure 1b: School enrolment, single mothers (1995-96 cut-off, age 3)



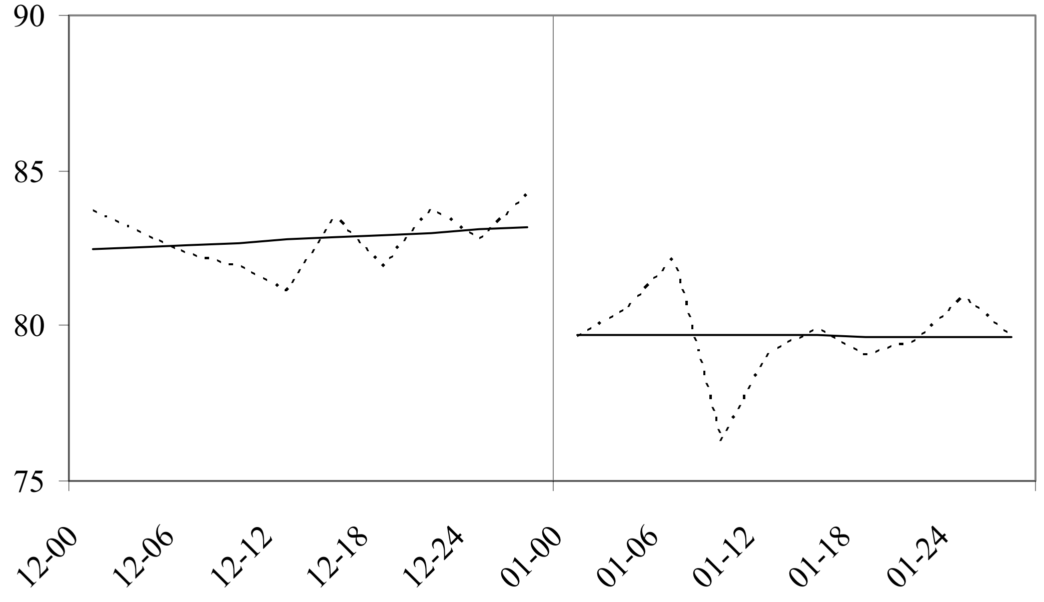
Source: 1999 census.
Sample: Single-parent families whose youngest child was born between Dec. 1995 and Jan. 1996.

Figure 2a : LM Participation, two-parent families (1995-96 cut-off, age 3)

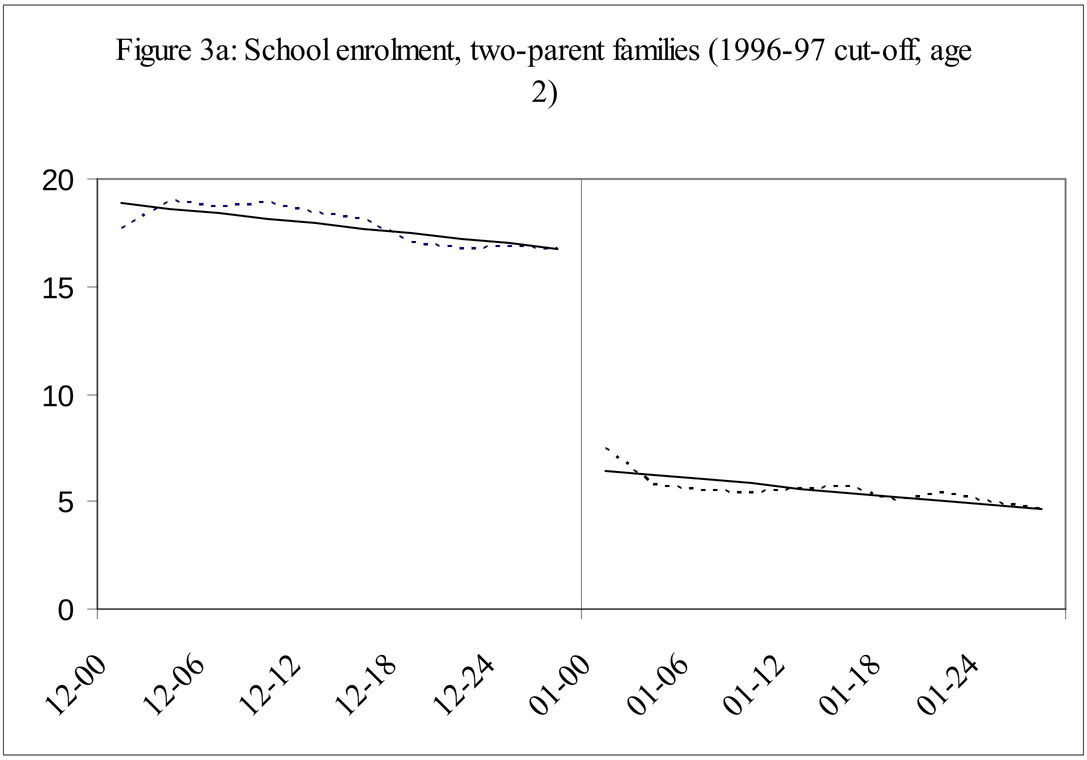


Source: 1999 census.
Sample: Two-parent families whose youngest child was born between Dec. 1995 and Jan. 1996.
Reading: The proportion of mothers participating in the LM is about 78% for mothers whose youngest child was born in Dec. 1995 as well as for those whose youngest child was born in Jan. 1996.

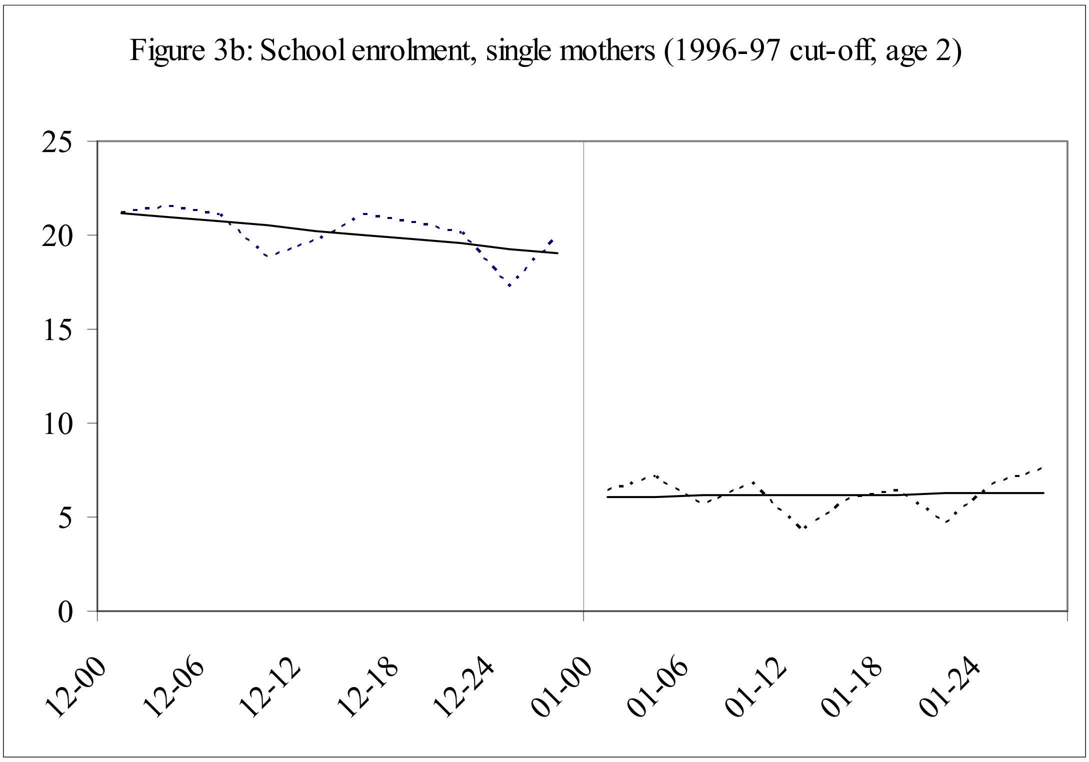
Figure 2b: LM Participation, single mothers (1995-96 cut-off, age 3)



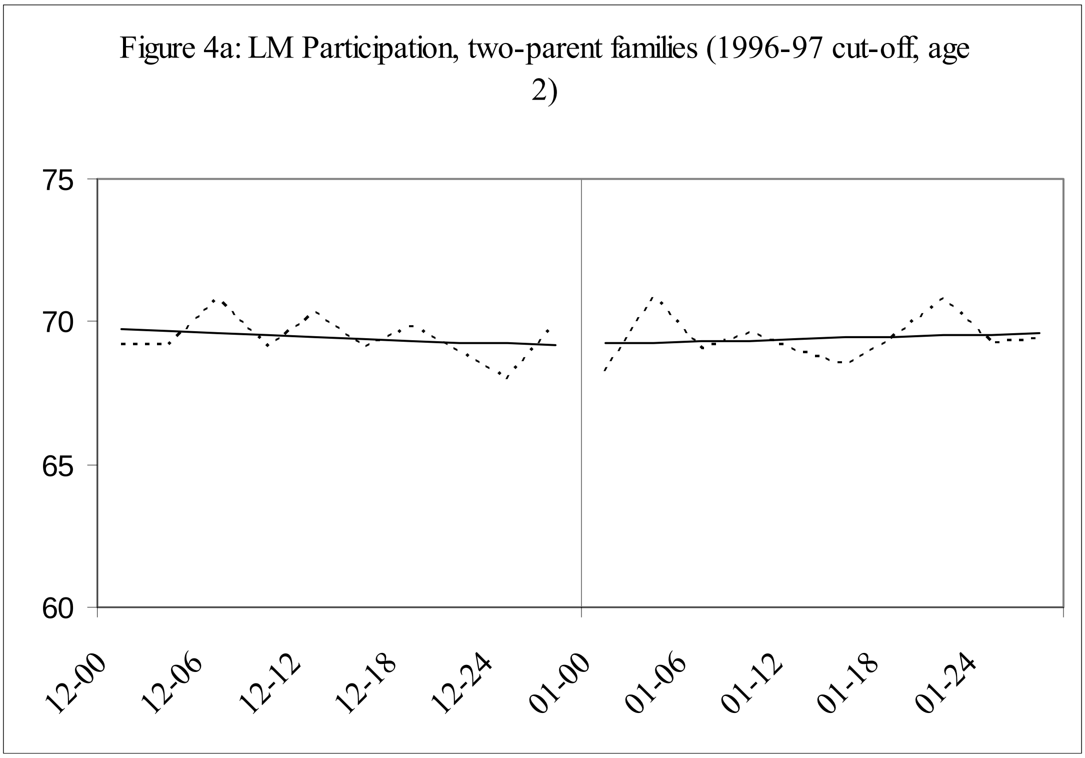
Source: 1999 census.
Sample: Single-parent families whose youngest child was born between Dec. 1995 and Jan. 1996.



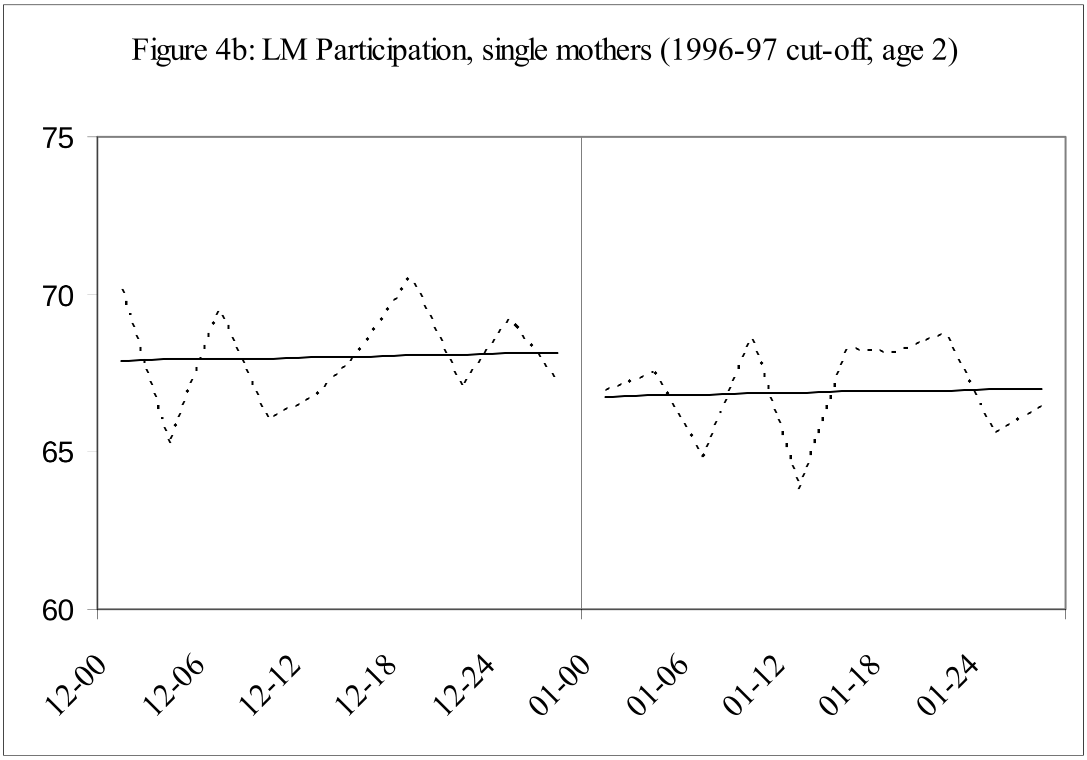
Source: 1999 census.
Sample: Two-parent families whose youngest child was born between Dec. 1996 and Jan. 1997.



Source: 1999 census.
Sample: Single-parent families whose youngest child was born between Dec. 1996 and Jan. 1997.

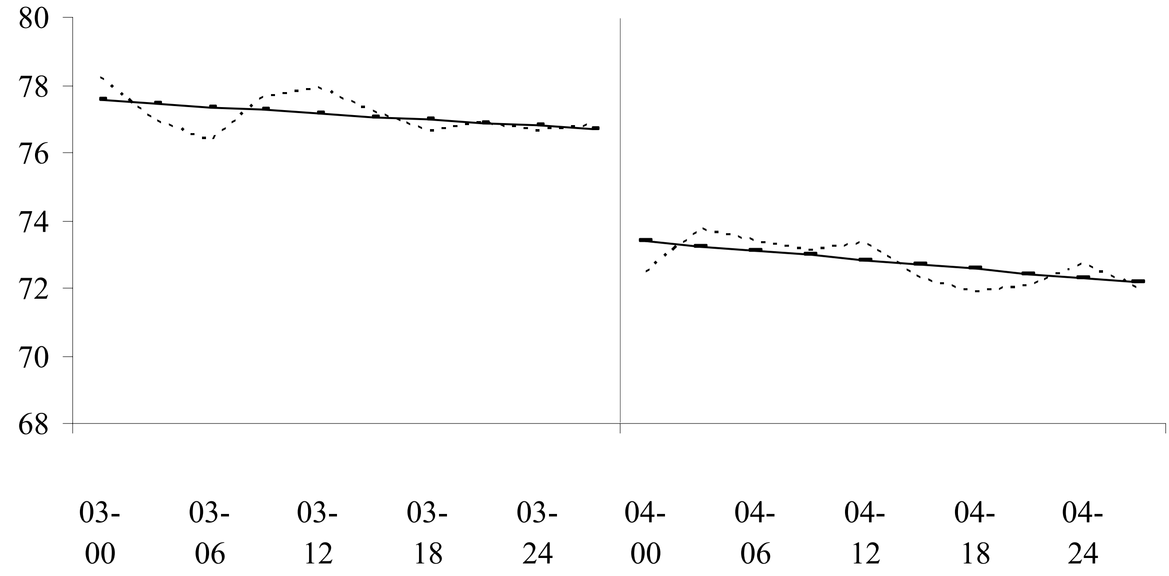


Source: 1999 census.
Sample: Two-parent families whose youngest child was born between Dec. 1996 and Jan. 1997.



Source: 1999 census.
Sample: Single-parent families whose youngest child was born between Dec. 1996 and Jan. 1997.

Figure 5: LM participation, two-parent families, (1996 March/April family benefits cutoff)

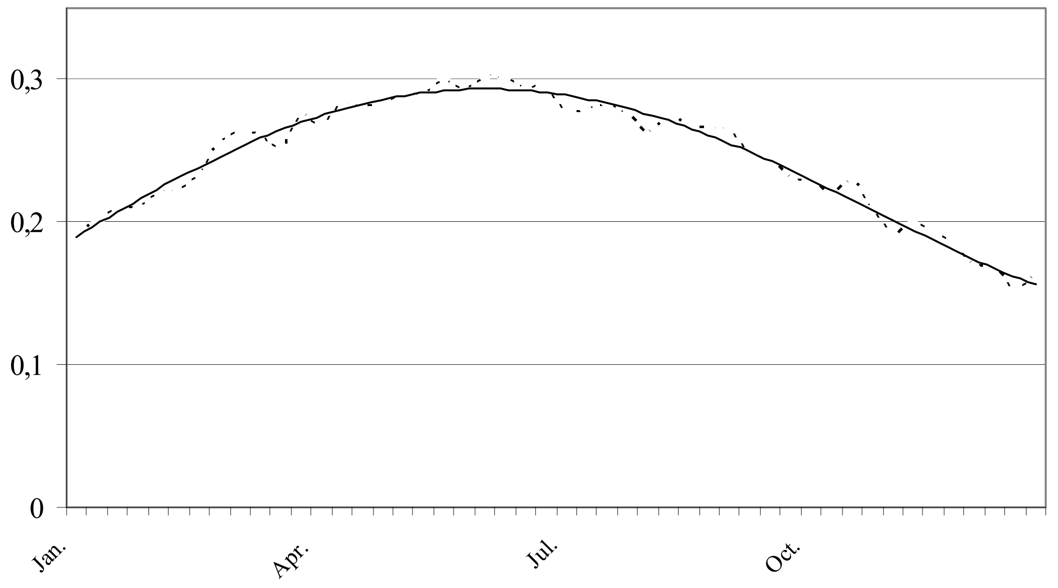


Source: 1999 census.
Sample: Two-parent families whose youngest child was born between March 1996 and April 1996.

Figure 6a: Difference in Early Preschool Enrolment Rates Across High and Low-Enrolment Regions, by Day of Birth, census 1999

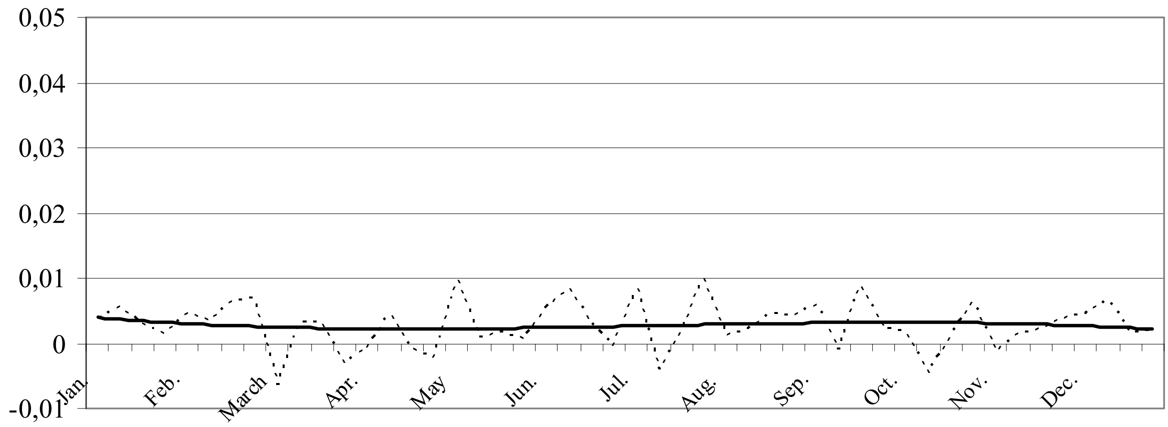
Source: 1999 census.

Figure 6b: Difference in Early Preschool Enrolment Rates Across High and Low-Enrolment Regions, by Day of Birth, census 1982.



Source: 1982 census.

Figure 7: Difference in Early Dropout Rate Across High and Low Enrolment Regions, by Day of Birth.



Source: 1999 census.

Table 1a: The Effect of Pre-elementary School Enrolment on Mothers' Labor Supply: Regression-Discontinuity Estimates, by Age Group and Family Type.

Panel A	Two-parent families			
	Children born between Dec. 1995, 1 and 1996, January, 31		Children born between Dec. 1996, 1 and 1997, January, 31	
	School Enrolment (first stage)	Mother's Participation (reduced form)	School Enrolment (first stage)	Mother's Participation (reduced form)
Intercept	0.938 (0.023)	0.754 (0.026)	0.262 (0.018)	0.709 (0.026)
Year of birth=1996	-0.171 (0.005)	-0.004 (0.006)	0.101 (0.004)	0.001 (0.006)
Day of birth	-0.051 (0.022)	0.028 (0.024)	-0.078 (0.017)	-0.014 (0.025)
Day of birth ×1996×100	-0.075 (0.030)	-0.013 (0.034)	-	-
Day of birth ×1997×100	-	-	0.014 (0.024)	0.029 (0.035)
R-Square	0.062	0.000	0.037	0.000
Number of observations	73,493	73,493	88,843	88,843
RD Estimate *Z=[Year of birth=1996]*	0.025 (0.036)		0.005 (0.062)	
OLS Estimate	0.085* (0.004)		-0.024 (0.005)	

Source: 1999 Census.
Sample: Families whose youngest child was born between Dec. 1995 and Jan. 1996 (two first columns) or between Dec. 1996 and Jan. 1997 (two last columns).

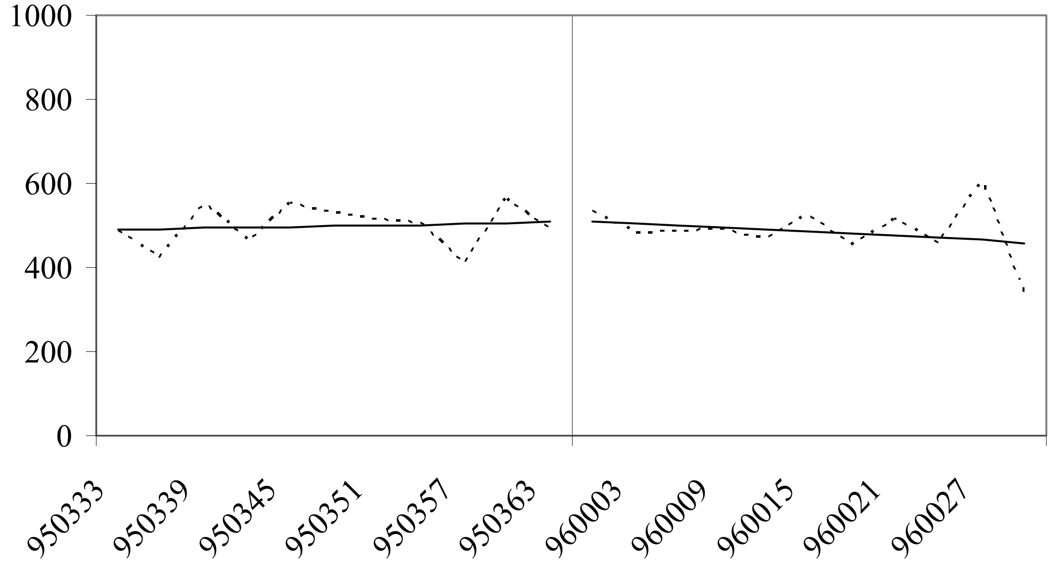
Table 1b: The Effect of Pre-elementary School Enrolment on Mothers' Labor Supply: Regression-Discontinuity Estimates, by Age Group and Family Type.

Panel B	Single-mother families			
	Children born between Dec. 1995, 1 and 1996, January, 31		Children born between Dec. 1996, 1 and 1997, January, 31	
	School enrolment (first stage)	Mother's Participation (reduced form)	School enrolment (first stage)	Mother's Participation (reduced form)
Intercept	0.905 (0.065)	0.797 (0.066)	0.285 (0.055)	0.668 (0.078)
Year of birth=1996	-0.143 (0.015)	-0.036 (0.015)	0.131 (0.013)	0.014 (0.018)
Day of birth	-0.016 (0.060)	0.029 (0.061)	-0.078 (0.052)	0.011 (0.073)
Day of birth ×1996×100	-0.110 (0.086)	-0.026 (0.086)	-	-
Day of birth ×1997×100	-	-	0.095 (0.073)	0.002 (0.103)
R-Square	0.044	0.002	0.043	0.000
Number of observations	10,188	10,188	10,471	10,471
RD Estimate *Z=[Year of birth=1996]*	0.253* (0.109)		0.111 (0.142)	
OLS Estimate	0.086* (0.010)		-0.011 (0.014)	

Source: 1999 Census.
Sample: Families whose youngest child was born between Dec. 1995 and Jan. 1996 (two first columns) or between Dec. 1996 and Jan. 1997 (two last columns).

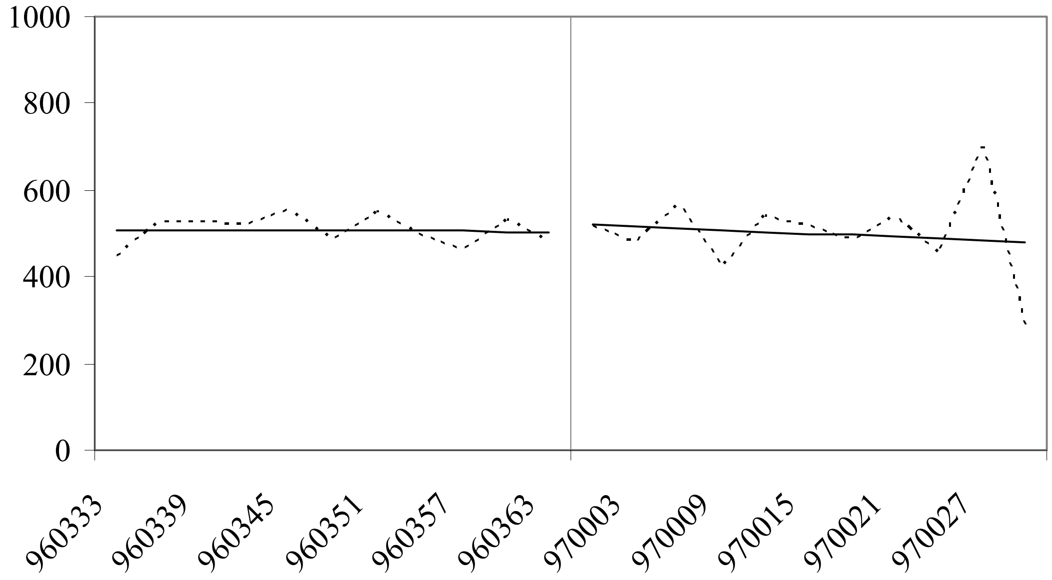
Appendix

Figure A1 : Number of Single-Parent Families, By Youngest Child Date of Birth (95-96 cutoff)



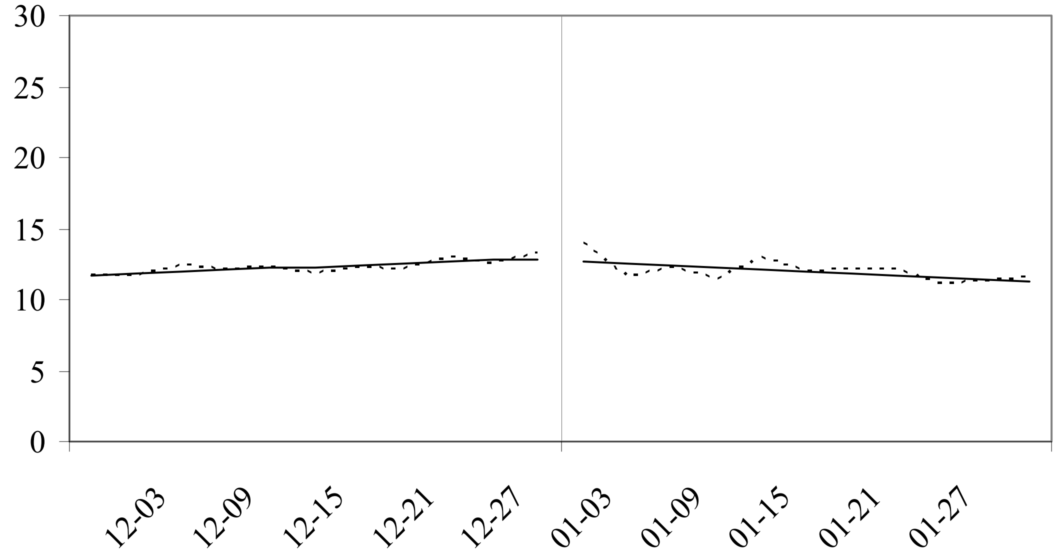
Source: 1999 census.
Sample: Single-parent families whose youngest child was born between December 1995 and January 1996.

Figure A2 : Number of Single-Parent Families, By Youngest Child Date of Birth (96-97 cutoff)



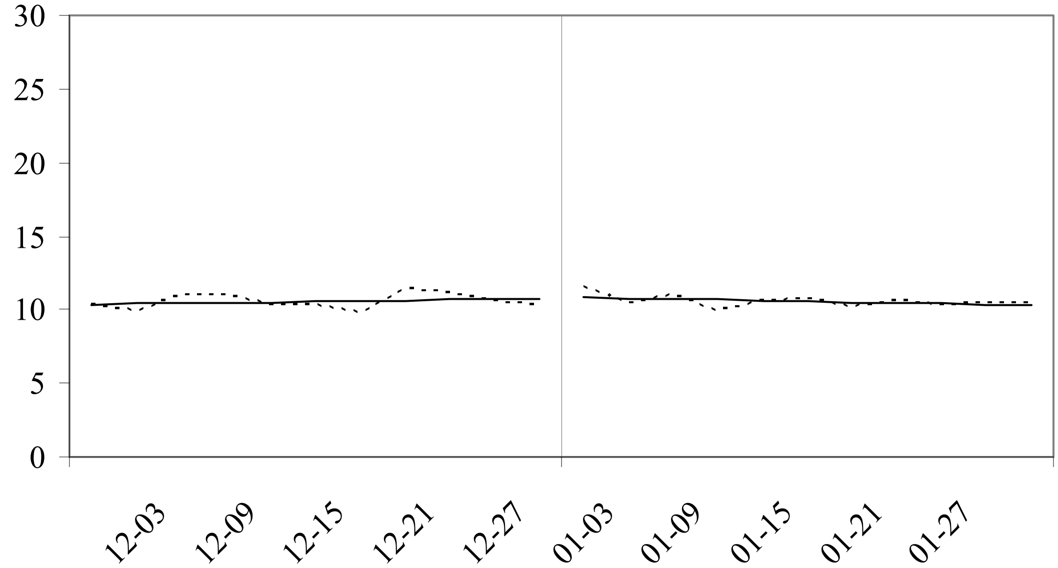
Source: 1999 census.
Sample: Single-parent families whose youngest child was born between December 1996 and January 1997.

Figure A3: Proportion of Single-Parent Families Among Families with One Child Born Between Dec. 1995 and January 1996.



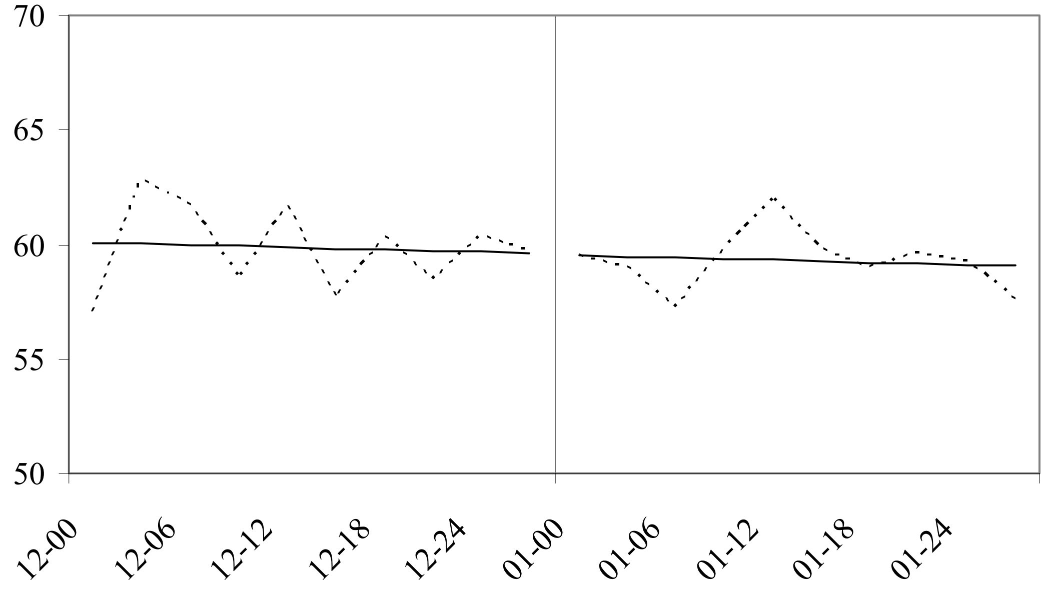
Source: 1999 census.
Sample: Families with one child born between December 1995 and January 1996.

Figure A4: Proportion of Single-Parent Families Among Families with One Child Born Between Dec. 1996 and January 1997.



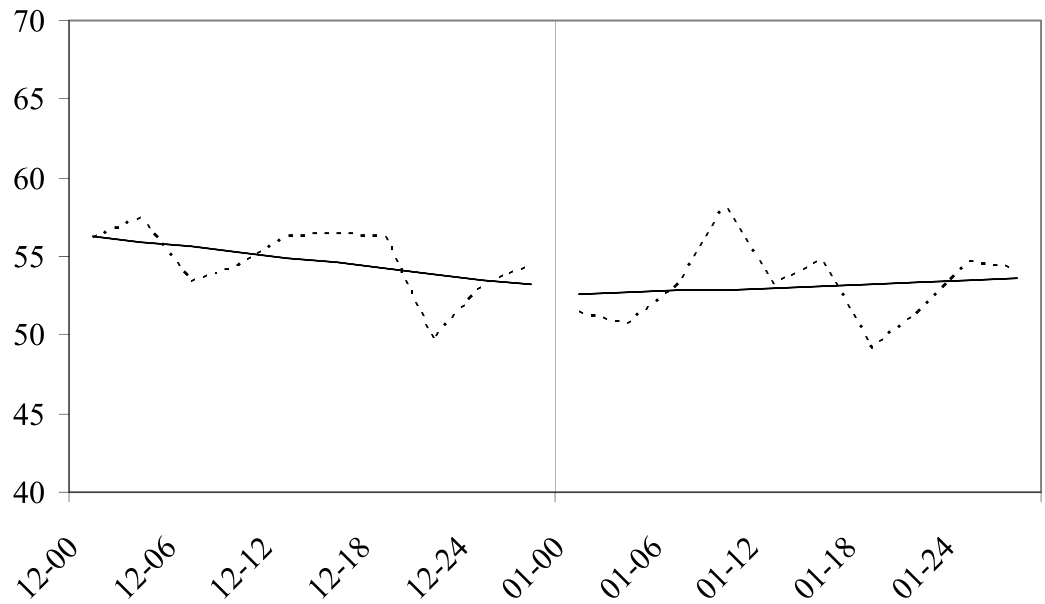
Source: 1999 census.
Sample: Families with one child born between December 1996 and January 1997.

Figure A5: Subsequent Fertility, Single-Parent Families (1995-96 cut-off).



Source: 1999 census.
Sample: Families with one child born between December 1995 and January 1996.

Figure A6 :Subsequent Fertility, Single-Parent Families (96-97 cut-off)



Source: 1999 census.
Sample: Families with one child born between December 1996 and January 1997.

Table A1: Estimates of Single-Mother Labor Supply Elasticity : Alternative Specifications.

	Children born between the 1st of Oct. 1995 and the 31st of March 1996		
	(1)	(2)	(3)
RD Estimate	.48*	.27*	.35*
	(.21)	(.10)	(.12)
	Children born between the 1st of Oct. 1996 and the 31st of March 1997		
	(1)	(2)	(3)
RD Estimate	.22*	.09	.19
	(.09)	(.12)	(.12)

Source : census 1999.

Sample: Families whose youngest child was born between October 1995 and March 1996 (three first rows) or between October 1996 and March 1997 (three last rows). Column (1) uses polynomials of degree 1 in exact date of birth as control function whereas column (2) uses polynomials of degree (2) and column (3) uses polynomials of degree 3.