Public Schooling for Young Children and Maternal Labor Supply

By JONAH B. GELBACH*

As policy makers have sought to reduce the welfare rolls by increasing labor supply among single mothers, much attention has been given to the possible role of expanded child care subsidies, including direct provision of public preschool. At the same time, public interest in child care subsidies for two-parent households is high; for example, former Vice President Gore not long ago proposed to make "high-quality pre-school fully available to every family, for every child, in every community in America" (Albert Gore, 1999).

In this paper, I use 1980 Census data to estimate the effect of public school enrollment for a woman's five-year-old on measures of labor supply and public assistance receipt. This approach has two advantages. First, I am able to estimate the effect of a large implicit child care subsidy.[1] Second, because public kindergarten is universally available to all age-eligible children where it is provided, selection problems related to means-testing do not arise.

A difficulty with public school enrollment arises because parents may choose to hold their children back a year or enroll them in private school. I deal with this issue by using five-year-old's quarter of birth (QOB) variables as instruments for public school enrollment status. This strategy works because parents' ability to enroll a child in public kindergarten in the academic year when the child turns five typically depends on the calendar date of the child's birth.[2] In most states, children born in the second quarter (April 1–June 30) of 1974 will have been eligible to start kindergarten in the fall of 1979, while children born in the first quarter (January 1–March 31) of 1975 generally will not. Eligibility for children born between July 1 and December 31, 1974 depends on the rules where they live.[3]

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[1] According to Craig Coelen et al. (1979), the nationwide average weekly fee paid by parents for full-time care in child care centers in 1976 ranged from $19–$26; inflating to 1979 prices yields a range of $24–$33. Children in such care averaged 40 hours of care per week, while 15 hours is a conservative estimate of the amount of care provided per week by kindergarten. Hence the average value of free public kindergarten in 1979 dollars was around $10–$13. Average 1979 weekly wage and salary income among single mothers with a five-year-old was just $80, so the implicit subsidy is certainly large for these women.

[2] QOB variables were first used as instruments by Joshua D. Angrist and Alan B. Krueger (1991) in their study examining the return to schooling. One criticism of this approach concerns omitted variables (see John Bound et al., 1995). I address this issue via overidentification testing. Bound et al. (1995) and Douglas Staiger and James H. Stock (1997) also raise statistical concerns related to the weak first-stage relationship between QOB variables and educational attainment in Angrist and Krueger's data. This concern does not arise here, because the first-stage relationship between QOB and public school enrollment of five-year-olds is very strong.

[3] Age-at-entry rules vary across states [data are available from U.S. Department of Education (various years) and details are also provided in William D. Valente (1985) and Angrist and Krueger (1991)]. However, unobservable determinants of both enrollment and labor-market outcomes may also vary systematically across states, so state dummy variables are unlikely to make good instruments. Using interactions of state and QOB dummies as instruments added little explanatory power to the first stage and did not appreciably affect my IV estimates. Adding the interactions while removing the QOB main effects from the instrument set led to very unstable estimates, because differences across states in the reduced-form effects of QOB on labor-market outcomes are slight.

My results provide significant evidence that public school enrollment has a substantial effect on maternal labor supply. Among single women whose youngest child is five, free public schooling for the five-year-old increased labor-supply measures by between 6–24 percent while reducing public assistance receipt by 10 percent. Among single mothers with both a five-year-old and a younger child, the estimates imply no significant impact of public school enrollment for the five-year-old. Among married mothers of five-year-olds, public school enrollment yields 6–15-percent increases in labor-supply measures, with little variation according to whether the woman also has a child younger than five.

The public schooling approach differs markedly from previous research on child care and female labor supply. Much attention has been paid to estimating structural parameters of utility functions, which are then used to simulate both labor-supply elasticities and the effects of various policy reforms. Child care cost variables generally are constructed from either observed household expenditures or area-level averages of prices or expenditures.[4] Examples include David M. Blau and Philip K. Robins (1988), Rachel Connelly (1992), Charles Michalopolous et al. (1992), and Jean Kimmel (1998). Most studies have found evidence of a significant negative labor-supply response to child care prices among married mothers, though the range of estimated employment elasticities is rather large, from 0 to −1.6. For single mothers, the literature yields no clear conclusion that child care costs have a significant impact on labor supply. For example, Kimmel (1998) reports elasticities ranging from −4.54 to +1.38.

Several explanations for the variation in results have been advanced, including quality heterogeneity (Blau and Alison P. Hagy, 1998), budget set nonlinearities and misspecification (Susan L. Averett et al., 1997), and simultaneity due to the use of area-level variation in child care costs (Gelbach, 1999). My focus on public school enrollment does not so much solve these problems as sidestep them. Providing access to public school can be thought of as offering a 100-percent marginal price subsidy for child care of fixed quality, as long as total consumption of nonfamily child care is less than the length of the school day. When total consumption of nonfamily child care exceeds the length of the school day, the subsidy is entirely inframarginal with respect to child care costs. Such a subsidy induces a kink in the budget constraint and so necessarily provides both price and income subsidies (e.g., see Gary Burtless and Jerry A. Hausman, 1978). As a result, my estimates are mixtures of price and income elasticities. Nonetheless, these estimates address the question of whether large subsidies in the form of limited, directly provided care influence maternal labor supply. This question is of clear interest to both labor economists and policy makers.

I. Public School Enrollment and Characteristics of the Sample

A. *Enrollment and QOB*

I begin by establishing a relationship between five-year-old's QOB and public school enrollment in my sample. Mothers who have both a five-year-old and a younger child must find child care for the younger child even if the five-year-old is enrolled, so throughout the paper I split the samples according to whether the woman's five-year-old is her youngest child. Table 1 reports public school enrollment rates by QOB for single and married mothers of five-year-olds in the 1980 Census.[5] For both single and married mothers, five-year-old's QOB is strongly correlated with public school enrollment, and public school enrollment rates are monotonically decreasing in QOB. Differences across quarters are highly statistically sig-

[4] A notable exception is Mark C. Berger and Dan A. Black (1992), who use women on the waiting list as a comparison group for recipients of child care subsidies in Kentucky.

[5] I use 1980 data because the 1990 Census did not ask respondents for QOB. The reference date for the Census is April 1, so data on age and year of birth together would allow me to distinguish first-quarter births from other births. While the 1990 long form did ask for both age and year of birth, the year of birth data are not included in the Public Use Microdata Sample. The construction of my sample is described in detail in Appendix A of my working paper (Gelbach, 1999).

TABLE 1—PUBLIC AND PRIVATE ENROLLMENT RATES BY QOB

		Quarter of Five-Year-Old's Birth				
Variable	Full Sample	74:II	74:III	74:IV	75:I	F
		Single Mothers				
Youngest is five years old						
Public school	0.63	0.83	0.79	0.55	0.31	807
Private school	0.16	0.11	0.13	0.18	0.22	46
N	10,932	2,646	3,044	2,715	2,527	
Youngest is younger than five						
Public school	0.65	0.86	0.81	0.58	0.32	554
Private school	0.08	0.06	0.07	0.10	0.11	15
N	6,885	1,718	1,910	1,649	1,608	
		Married Mothers				
Youngest is five years old						
Public school	0.55	0.83	0.75	0.43	0.18	6,374
Private school	0.21	0.13	0.16	0.25	0.31	517
N	53,163	12,784	14,245	13,581	12,553	
Youngest is younger than five						
Public school	0.54	0.81	0.75	0.40	0.16	6,644
Private school	0.19	0.13	0.15	0.23	0.26	312
N	52,134	12,908	14,309	13,049	11,868	

Notes: Number in last column is F-statistic for test of equivalence across quarters. Standard errors for all coefficients were 0 to two decimal places, as were p-values for all F-statistics.

nificant; for example, the F-statistic testing equivalence of public school enrollment across quarters among single mothers whose youngest child is five is 807, yielding a p-value of 0. Hence QOB dummies in this context are very strong instruments and are not subject to the weak first-stage critiques lodged by Bound et al. (1995) and Staiger and Stock (1997).

Private school enrollment rates move monotonically in the opposite direction to public school enrollment. For example, only 11 percent of 74:Q2 births to single mothers whose youngest child is five are enrolled in private school, while 22 percent of 75:Q1 births are enrolled. These figures suggest that a substantial fraction of parents of five-year-olds substitute public school for private school when possible.

B. *Demographic Characteristics*

Tables 2 and 3 show means of demographic covariates I use in the estimation below. The figures in the first column of Table 2 refer to the entire sample of single mothers and suggest that those whose youngest child is five differ significantly from those who also have a younger child. Those who also have a younger child have nearly a year less education, are significantly more likely to live in central cities, and are significantly less likely to be black. They are also several years younger and have significantly fewer older children, differences that should be expected since the presence of both a five-year-old and a younger child indicates a different stage of the life cycle than the presence of only a five-year-old.

For single mothers whose youngest child is five, significant differences across five-year-old's QOB appear only for mother's age variables and the number of children aged 13–17. These differences are mechanically related to the fact that mothers of children born in later quarters will tend to be older themselves. For single mothers whose youngest child is younger than five, significant differences across QOB also appear for mother's educational attainment, central city residence, and race. Simulations suggested that these differences can be accounted for largely by racial differences in birth spacing; because of differences in the size of each subpopulation, these differences need not affect the sample of single mothers whose youngest child is five.[6] Since demographic variables are statistically associated with both QOB and labor supply, I control for them in the IV estimation.

[6] These issues are discussed in detail in Appendix C of Gelbach (1999).

TABLE 2—DEMOGRAPHIC CHARACTERISTICS BY QOB AND YOUNGEST CHILD'S AGE, SINGLE MOTHERS

Variable	Full Sample	Quarter of Five-Year-Old's Birth 74:II	74:III	74:IV	75:I	F
Youngest Is Five Years Old						
Age of mother	30.7	31.1	30.8	30.7	30.2	10.65
Squared age of mother	974	997	979	974	945	8.67
Mother's years of education	11.72	11.70	11.75	11.74	11.65	0.82
Central city	0.32	0.33	0.31	0.32	0.32	1.25
White	0.62	0.61	0.63	0.63	0.62	0.42
Number of children in household						
0–5	1.00	1.00	1.00	1.00	1.00	
6–12	0.79	0.82	0.79	0.78	0.79	0.78
13–17	0.27	0.30	0.28	0.27	0.23	5.20
≥18	0.07	0.07	0.08	0.06	0.06	1.19
Number of other household members						
<18	0.07	0.08	0.07	0.06	0.07	0.94
≥18	0.20	0.20	0.20	0.19	0.20	0.25
N	10,932	2,646	3,044	2,715	2,527	
Youngest Is Younger Than Five						
Age of mother	27.6	27.8	27.7	27.7	27.2	5.40
Squared age of mother	786	792	793	794	764	3.87
Mother's years of education	10.92	11.08	10.94	10.83	10.83	3.68
Central city	0.38	0.36	0.37	0.37	0.41	4.15
White	0.49	0.52	0.48	0.49	0.46	4.40
Number of children in household						
0–5	2.27	2.29	2.27	2.26	2.25	1.86
6–12	0.63	0.57	0.63	0.65	0.66	3.66
13–17	0.13	0.12	0.15	0.14	0.12	1.53
≥18	0.03	0.03	0.03	0.03	0.03	0.08
Number of other household members						
<18	0.07	0.05	0.06	0.07	0.09	2.67
≥18	0.19	0.19	0.19	0.19	0.19	0.02
N	6,885	1,718	1,910	1,649	1,608	

Notes: Top number in last column is F-statistic for test of equivalence across quarters; number in parentheses is associated p-value. p-values (critical values) of F-statistics are: 0.01 (3.78), 0.05 (2.61), and 0.10 (2.08).

However, exclusion of the demographic variables does not alter the IV estimates among single mothers whose youngest is five. For single mothers whose youngest is younger than five, excluding the demographic controls changes the point estimates for public school enrollment but does not alter the substantive conclusions drawn here.

Demographic characteristics for married mothers are reported in Table 3. On average, married mothers in my sample are about a year and a half older than the single mothers, have somewhat more education, are substantially less likely to live in central cities, and are much more likely to be white. Differences by youngest child's age in the age of the mother and her children's age structure mirror those for single mothers. Lastly, despite the enormous sample sizes among married mothers, there are significant differences across QOB in only four other instances (central city residence; number of own children 13–17 among married mothers whose youngest child is five; number of children aged 0–5; and number of other household members younger than 18 among married mothers whose youngest child is younger than five).

II. Results

The kink in the budget constraint caused by public school enrollment means that some women who enroll their five-year-old receive a marginal price subsidy, some receive only an income subsidy, and some (those at the kink) receive a mixture of the two. Intuition suggests that estimates of the impact of public school enrollment should be a weighted average of

TABLE 3—DEMOGRAPHIC CHARACTERISTICS BY QOB AND YOUNGEST CHILD'S AGE, MARRIED MOTHERS

		Quarter of Five-Year-Old's Birth				
Variable	Full Sample	74:II	74:III	74:IV	75:I	F
		Youngest Is Five Years Old				
Age of mother	32.7	33.1	32.8	32.7	32.3	45.55
Squared age of mother	1,098	1,121	1,104	1,095	1,072	40.93
Mother's years of education	12.27	12.27	12.24	12.29	12.29	1.20
Central city	0.15	0.14	0.15	0.16	0.15	3.31
White	0.86	0.87	0.86	0.86	0.86	1.77
Number of children in household						
0–5	1.00	1.00	1.00	1.00	1.00	
6–12	0.96	0.95	0.96	0.97	0.96	1.82
13–17	0.30	0.32	0.30	0.30	0.28	8.49
≥18	0.07	0.07	0.07	0.07	0.06	1.32
Number of other household members						
<18	1.03	1.03	1.02	1.03	1.03	1.59
≥18	0.06	0.06	0.06	0.06	0.06	0.30
N	53,163	12,784	14,245	13,581	12,553	
		Youngest Is Younger Than Five				
Age of mother	29.4	29.7	29.5	29.3	29.0	52.43
Squared age of mother	884	903	888	879	863	46.96
Mother's years of education	12.26	12.29	12.27	12.29	12.21	1.89
Central city	0.16	0.16	0.16	0.16	0.15	2.04
White	0.85	0.85	0.85	0.85	0.85	1.16
Number of children in household						
0–5	2.24	2.26	2.25	2.23	2.22	19.21
6–12	0.51	0.52	0.50	0.51	0.52	1.98
13–17	0.08	0.08	0.08	0.07	0.08	1.21
≥18	0.01	0.01	0.01	0.01	0.01	0.13
Number of other household members						
<18	1.02	1.03	1.02	1.02	1.02	4.57
≥18	0.06	0.07	0.06	0.06	0.06	0.88
N	52,134	12,908	30,056	13,049	11,868	

Notes: Top number in last column is F-statistic for test of equivalence across quarters; number in parentheses is associated p-value. p-values (critical values) of F-statistics are: 0.01 (3.78), 0.05 (2.61), and 0.10 (2.08).

effective price and income effects. As I show in Appendix B of my working paper (Gelbach, 1999), as long as QOB is independent of the enrollment effect itself, this intuition is correct. Subject to instrument validity, IV estimates nonparametrically identify the average effect of public schooling, where the weights on the price and income effects are the fraction of the population that actually receives the relevant price and/or income subsidy. It is important to keep in mind that if there is heterogeneity in the enrollment effects, only the effect of treatment on women observed enrolling their children is identified.[7]

I focus on seven outcomes: weeks of work in 1979; employment in 1979; employment last week; usual weekly hours in 1979; hours of work last week; public assistance receipt in 1979; and wage and salary income in 1979. Because a child would have had to enroll at the beginning of the school year (September 1979) in order to be enrolled in most public school programs as of the 1980 Census, the enrollment variable is appropriate for analyzing both 1979 annual and 1980 last-week outcome measures.[8]

A. *Outcomes and QOB*

In the first column of Table 4, I report average values of these outcome levels. Average

[7] These conclusions follow from well-known results in the literature on evaluation methodology (e.g., James J. Heckman and Richard Robb, Jr., 1985).

[8] Even an eligible five-year-old in 1980 generally will have had access to public schooling only for the months of September through December, so estimates using 1979 labor-supply variables will underestimate the effects of full-year free child care.

TABLE 4—OUTCOMES BY QOB AND YOUNGEST CHILD'S AGE, SINGLE MOTHERS

		Quarter of Five-Year-Old's Birth				
Variable	Full Sample	74:II	74:III	74:IV	75:I	F
	Youngest Is Five Years Old					
Weeks of work in 1979	28.36	29.60	28.28	28.45	27.07	6.369
	(0.22)	(0.40)	(0.38)	(0.40)	(0.41)	(0.000)
Employment in 1979	0.700	0.715	0.701	0.695	0.689	1.855
	(0.004)	(0.008)	(0.008)	(0.008)	(0.008)	(0.135)
Usual weekly hours in 1979	25.7	26.4	25.9	25.4	25.1	2.687
	(0.2)	(0.3)	(0.3)	(0.3)	(0.3)	(0.045)
Employment last week	0.589	0.607	0.586	0.593	0.569	3.077
	(0.005)	(0.009)	(0.008)	(0.009)	(0.009)	(0.026)
Hours of work last week	21.5	22.2	21.8	21.4	20.6	3.325
	(0.2)	(0.4)	(0.3)	(0.4)	(0.4)	(0.019)
1979 receipt of public assistance	0.343	0.325	0.343	0.353	0.353	2.520
	(0.005)	(0.008)	(0.008)	(0.008)	(0.009)	(0.056)
1979 wage and salary income	5,193	5,483	5,276	5,050	4,946	5.074
	(56)	(106)	(99)	(104)	(108)	(0.002)
	Youngest Is Younger Than Five					
Weeks of work in 1979	15.95	16.85	15.09	16.02	15.94	2.578
	(0.25)	(0.46)	(0.44)	(0.47)	(0.48)	(0.052)
Employment in 1979	0.488	0.509	0.474	0.484	0.485	1.800
	(0.006)	(0.011)	(0.011)	(0.011)	(0.011)	(0.145)
Usual weekly hours in 1979	17.4	18.1	16.7	17.5	17.3	1.697
	(0.2)	(0.4)	(0.4)	(0.5)	(0.5)	(0.165)
Employment last week	0.341	0.355	0.316	0.352	0.345	3.099
	(0.006)	(0.011)	(0.010)	(0.011)	(0.011)	(0.026)
Hours of work last week	11.7	12.0	11.0	12.1	11.8	1.595
	(0.2)	(0.4)	(0.4)	(0.4)	(0.4)	(0.188)
1979 receipt of public assistance	0.551	0.555	0.553	0.546	0.549	0.122
	(0.006)	(0.011)	(0.011)	(0.012)	(0.011)	(0.947)
1979 wage and salary income	2,525	2,544	2,482	2,509	2,574	0.152
	(54)	(103)	(98)	(105)	(107)	(0.929)

Notes: White-robust standard errors are in parentheses. Top number in last column is F-statistic for test of equivalence across quarters; number in parentheses is associated p-value. Reported figures in first five columns are coefficients on a full set of QOB dummies in regressions including state fixed effects and demographic variables listed in Table 2. All state fixed effects and demographic control variables are demeaned, so that QOB coefficients represent the conditional mean outcomes for the relevant birth quarter when all other variables are held at their means.

outcome levels differ greatly according to whether the woman has a child younger than five. For example, single mothers with both a five-year-old and a younger child worked an average of only 16 weeks in 1979, by comparison to 28 weeks among single mothers whose youngest is five. Similarly, 55 percent of single mothers with a child younger than five received public assistance in 1979, by comparison to only 34 percent of those whose youngest is five. While child care needs of children younger than five can be expected to cause some differences in outcomes across the populations, the magnitude of the differences seems too large to attribute only to this difference. A complementary explanation is that the choice to have another child between the five-year-old's birth and 1980 is correlated with other variables that are directly related to labor-market outcomes.

The next four columns of Table 4 report average values of the outcome variables across five-year-old's QOB among single mothers. These figures were computed by regressing the outcomes on four QOB dummies (with no constant included), fixed effects for state of residence and five-year-old's state of birth, and all demographic variables listed in Table 2. All non-QOB variables in these regressions were demeaned, so the figures represent average values of the outcomes by QOB when all other

TABLE 5—OUTCOMES BY QOB AND YOUNGEST CHILD'S AGE, MARRIED MOTHERS

Variable	Full Sample	Quarter of Five-Year-Old's Birth				F
		74:II	74:III	74:IV	75:I	
Youngest Is Five Years Old						
Weeks of work in 1979	21.72	22.05	22.06	21.69	21.03	6.402
	(0.10)	(0.19)	(0.18)	(0.19)	(0.19)	(0.000)
Employment in 1979	0.578	0.589	0.585	0.578	0.558	10.778
	(0.002)	(0.004)	(0.004)	(0.004)	(0.004)	(0.000)
Usual weekly hours in 1979	19.1	19.5	19.4	19.1	18.4	9.458
	(0.1)	(0.2)	(0.2)	(0.2)	(0.2)	(0.000)
Employment last week	0.483	0.493	0.494	0.480	0.463	11.260
	(0.002)	(0.004)	(0.004)	(0.004)	(0.004)	(0.000)
Hours of work last week	15.6	15.9	15.9	15.6	15.0	7.639
	(0.1)	(0.2)	(0.2)	(0.2)	(0.2)	(0.000)
1979 receipt of public assistance	0.022	0.021	0.023	0.022	0.021	0.625
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.599)
1979 wage and salary income	3,532	3,536	3,599	3,525	3,460	1.679
	(23)	(45)	(42)	(43)	(45)	(0.169)
Youngest Is Younger Than Five						
Weeks of work in 1979	14.24	14.85	14.36	13.94	13.75	8.031
	(0.09)	(0.17)	(0.16)	(0.17)	(0.18)	(0.000)
Employment in 1979	0.438	0.452	0.439	0.434	0.427	6.214
	(0.002)	(0.004)	(0.004)	(0.004)	(0.004)	(0.000)
Usual weekly hours in 1979	13.8	14.2	13.9	13.6	13.4	4.857
	(0.1)	(0.2)	(0.1)	(0.2)	(0.2)	(0.002)
Employment last week	0.327	0.339	0.333	0.321	0.314	7.807
	(0.002)	(0.004)	(0.004)	(0.004)	(0.004)	(0.000)
Hours of work last week	9.9	10.1	10.2	9.7	9.4	6.996
	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)	(0.000)
1979 receipt of public assistance	0.030	0.030	0.030	0.029	0.030	0.065
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.979)
1979 wage and salary income	2,163	2,249	2,168	2,123	2,108	2.854
	(19)	(37)	(35)	(37)	(39)	(0.036)

Notes: White-robust standard errors are in parentheses. Top number in last column is F-statistic for test of equivalence across quarters; number in parentheses is associated p-value. Reported figures in first five columns are coefficients on a full set of QOB dummies in regressions including state fixed effects and demographic variables listed in Table 3. All state fixed effects and demographic control variables are demeaned, so that QOB coefficients represent the conditional mean outcomes for the relevant birth quarter when all other variables are held at their means.

included observables are held at their means. I report F-statistics testing the equivalence of outcomes across QOB in the last column of Table 4; significant F-statistics constitute evidence of a systematic reduced-form relationship between outcomes and QOB.

For single mothers whose youngest child is five, the results suggest a statistically significant relationship between QOB and six of the seven outcomes. The relationship between outcomes and QOB is essentially monotonic: labor-supply and earned income measures decrease as we move from earlier to later quarters of birth, while the frequency of public assistance receipt increases. For single mothers with both a five-year-old and a younger child, cross-QOB equivalence of outcomes can be rejected only for weeks of work in 1979 and employment last week. Given the significant relationship between public school enrollment and QOB in Table 1, these findings suggest that enrollment has no impact on outcomes among single mothers with both a five-year-old and a younger child.

Table 5 shows analogous figures for married mothers of five-year-olds. As with single mothers, labor-supply measures are significantly lower for mothers with both a five-year-old and

TABLE 6—ESTIMATES OF THE EFFECT OF PUBLIC SCHOOL ENROLLMENT ON HOURS WORKED LAST WEEK (YOUNGEST CHILD IS FIVE YEARS OLD)

Right-Hand-Side Variable	(1) Means	(2) OLS	(3) OLS	(4) 2SLS
Public school enrollment	0.632	−3.784	−1.009	2.707
of five-year-old	(0.005)	(0.400)	(0.383)	(0.885)
Number of own children in household aged:				
6–12	0.794		−3.753	−3.893
	(0.008)		(0.223)	(0.226)
13–17	0.270		−1.761	−1.909
	(0.006)		(0.320)	(0.323)
≥18	0.067		0.309	0.187
	(0.003)		(0.657)	(0.660)
Number of other household members aged:				
<18	0.069		−2.350	−2.392
	(0.003)		(0.518)	(0.521)
≥18	0.199		1.468	1.484
	(0.005)		(0.383)	(0.384)
Mother's years of	11.716		1.513	1.547
education	(0.025)		(0.075)	(0.075)
White	0.623		3.180	3.566
	(0.005)		(0.428)	(0.438)
Live in central city	0.320		−1.945	−1.862
	(0.004)		(0.435)	(0.437)
Age of mother	30.697		3.379	3.388
	(0.054)		(0.310)	(0.312)
Squared age of mother	974.224		−0.049	−0.049
	(3.562)		(0.005)	(0.005)

Note: Standard errors are in parentheses.

a younger child. Holding constant the presence of a child younger than five, labor supply is also lower among married mothers than among single mothers; these differences are relatively larger among women whose youngest child is five. Among both populations of married mothers, receipt of public assistance is nearly zero, suggesting that most public assistance reported in the sample of single mothers likely involves programs with categorical eligibility tests that exclude married couples.

B. *Outcomes and Enrollment*

As an example of the specifications to be estimated below, Table 6 presents estimates of the enrollment effect on hours of work last week among single mothers whose youngest child is five. Means of right-hand-side variables are shown in column (1). Column (2) offers a univariate ordinary least-squares (OLS) specification, with no controls included. The endogeneity and selection problems discussed earlier regarding the public school enrollment dummy are clearly displayed in this column: public school enrollment appears to have a *negative* effect on labor supply, being associated with a reduction in labor supply by 3.8 hours of work per week. Results in column (3) indicate that much of this negative relationship can be accounted for by demographic covariates, as the magnitude of the enrollment coefficient falls by almost 75 percent. These results reflect the correlation of five-year-old's public school enrollment with other observable variables that are associated with labor supply.

In column (4), I retain the controls from column (3) but instrument for public school enrollment using three QOB dummies. The estimated enrollment-induced increase of 2.7 hours of work is statistically very significant—more than three times its standard error. Coefficient estimates for the demographic variables in these equations are consistent with previous labor-supply studies and generally confirm predictions from economic theory.

The interpretation of the children's age variables is complicated by concerns that labor sup-

TABLE 7—EFFECT OF PUBLIC SCHOOL ENROLLMENT ON LABOR-SUPPLY MEASURES, SINGLE MOTHERS WHOSE YOUNGEST CHILD IS FIVE YEARS OLD

Left-Hand-Side Variable	(1) OLS	(2) OLS	(3) 2SLS	(4) Test of Overidentification
Employment in 1979	−0.076	−0.013	0.040	1.45
	(0.009)	(0.008)	(0.020)	(0.48)
Usual hours in 1979	−3.46	−0.92	2.23	0.82
	(0.38)	(0.36)	(0.83)	(0.66)
Weeks of work in 1979	−4.28	−1.15	3.60	5.63
	(0.45)	(0.42)	(0.99)	(0.06)
Received public assistance in 1979	0.095	0.034	−0.044	2.97
	(0.009)	(0.009)	(0.021)	(0.23)
Wage and salary income in 1979	−1,055.0	−430.1	931.6	2.04
	(116.2)	(111.2)	(258.9)	(0.36)
Hours last week	−3.78	−1.03	2.71	0.54
	(0.40)	(0.38)	(0.89)	(0.76)
Current employment	−0.085	−0.020	0.051	3.48
	(0.010)	(0.009)	(0.021)	(0.17)
State fixed effects	No	Yes	Yes	
Demographic controls	No	Yes	Yes	
Instruments	No	No	3 QOB dummies	

Notes: For columns (1)–(3), the top number in each cell is the coefficient on the public school enrollment dummy in equation estimated using the indicated row variable on the left-hand side. Standard errors are reported in parentheses. Column (4) top number is χ^2 overidentification test statistic with associated p-value in parentheses. Sample size is 10,932.

ply and family structure may be jointly chosen. However, taken at face value the results suggest that having another child aged 6–12 is associated with a significant reduction of 3.9 hours, which should be expected, since these children will require after-school child care. Adding an own child aged 13–17 to the woman's household is associated with a smaller reduction (1.9 hours), while the number of own children aged 18 or older in the household has little association with hours of work. Given the total number of own children, changes in age structure also are significantly associated with labor supply: having a 13–17-year-old instead of a 6–12-year-old is associated with an *increase* of 2.0 hours.

The relationship between hours of work and the number of other household members also depends on age structure. The number of other household members aged 18 or greater is associated with an increase of about 1.5 hours of work, while the number of other household members aged younger than 18 is associated with a reduction of about 2.4 hours of work. This relationship is to be expected, because older household members may be able to care for young children, while younger household members may add to household child care costs for women who wish to work.

The state of birth fixed effects were insignificant in the IV specification. However, state of residence fixed effects were jointly significant with a p-value of essentially 0. Since the fixed effects are estimated conditional on enrollment status, this finding raises doubts about the common practice of relying on area-level variation in child care prices to identify labor-supply equations.

Table 7 reports the coefficients on the public school enrollment dummy for all seven outcome measures among single mothers whose youngest child is five.[9] The pattern across specifications mirrors that seen in Table 6: the simple OLS estimate is wrong-signed and large in magnitude; including controls greatly reduces the estimated magnitude, but without changing its sign; and the IV estimates suggest a significant positive effect of enrollment on labor-supply measures.

A concern regarding instrument validity

[9] Full tables including the estimated coefficients for the demographic covariates are available on request.

arises because five-year-old's age may affect mother's labor-market outcomes directly. For example, Angrist and William Evans (1998) find that women aged 21–35 work less when they have a third child than when they do not, with the effect declining with the third child's age. Suppose that labor-market outcomes depend both on child care costs as measured by public school enrollment and on whether the five-year-old is older than five and one-half, i.e., was born after 74:Q3 or not. In this case, mean outcomes conditional on enrollment will be the same for mothers of 74:Q4 and 75:Q1 births. Thus with 75:Q1 as the excluded category, the Q4 dummy is a valid instrument. But conditional on enrollment, mean outcomes for 75:Q1 mothers will differ systematically from mean outcomes for 74:Q2 and 74:Q3 mothers, invalidating the Q2 and Q3 dummies. Since in this case the Q4 dummy is a valid instrument, overidentification testing will be informative.

In column (4) of Table 7, I report a standard Generalized Method of Moments (GMM)-based χ^2 overidentification test statistic (see William H. Greene, 1993) using the White-robust covariance matrix. The results show that the QOB instrumental variables easily pass the overidentification test for all but the 1979 weeks of work outcome variables.

To gain some idea of the magnitude of the public school enrollment effect, I construct a relative measure. Let $\bar{y}_e$ be the observed average level of outcome y among women who enrolled their five-year-old in public school. The estimated enrollment coefficient is the causal effect among enrollers, so the no-enrollment baseline is given by $\bar{y}_b = \bar{y}_e - \hat{\beta}$. A measure of the effect of enrollment relative to baseline labor supply is thus $R = \hat{\beta}/\bar{y}_b$.[10] The estimates imply that for single mothers whose youngest child is five, public schooling increased the probability of work in 1979 by 6 percent; increased usual hours of work in 1979 by 10 percent; increased weeks of work by 16 percent; reduced the probability of welfare receipt by 10 percent; increased wage and salary income by 24 percent; increased hours worked in the week before filling out the Census questionnaire by 16 percent; and increased current employment by 10 percent. In each case, the estimated percentage effect is significantly different from zero at the 0.05 level or lower.

Both employment and usual hours in 1979 appear to respond by more than for last week, though these differences are not significantly different. Since enrollment affects child care costs for only the latter part of 1979, we might expect the percentage effects for 1979 variables to be smaller than effects for their last-week counterparts. However, the effect of enrollment over several months depends on the joint distribution of the effects for all weeks during the period. If the enrollment effect on labor supply is serially correlated, then there is no clear way to compare the 1979 and last-week effects.

Also of interest is the fact that wage and salary income responds a good deal more than the labor-supply variables. One explanation is that women receiving child care through public schooling may be able to take jobs with longer hours, higher wages, or both. If there are fixed costs to working, or working full time, then the number of hours of child care provided by public schooling need not be large to induce the differences reported above. Alternatively, women who receive free child care through public schools may be more likely to report labor income. Various components of the welfare system impose large implicit taxes on labor income, so participants have an incentive not to report income to the government. It is well known that reported consumption among welfare participants in household surveys is greater than reported income, suggesting that women who earn income under the table may be hesitant to report that income to official survey-takers.[11] If access to free child care causes women not to participate in welfare

[10] I used the delta method to compute standard errors for these percentage effects. Ignoring covariance between estimated values of $\bar{y}_e$ and $\hat{\beta}$, the standard errors are $\sqrt{(\partial R/\partial \hat{\beta})^2 \hat{V}(\hat{\beta}) + (\partial R/\partial \bar{y}_e)^2 \hat{V}(\bar{y}_e)}$. In practice, the variance of $\bar{y}_e$ is negligible compared to the variance of $\hat{\beta}$, so the standard errors are approximately equal to $(\partial R/\partial \hat{\beta})\sqrt{(\hat{V}(\hat{\beta}))}$.

[11] Kathryn Edin and Laura Lein's (1997) informal survey of low-income women finds that nearly all welfare recipients have at least some labor income that they do not report to their caseworkers.

TABLE 8—EFFECT OF PUBLIC SCHOOL ENROLLMENT ON LABOR-SUPPLY MEASURES, SINGLE MOTHERS WHOSE YOUNGEST CHILD IS YOUNGER THAN FIVE YEARS OLD

Left-Hand-Side Variable	(1) OLS	(2) OLS	(3) 2SLS	(4) Test of Overidentification
Employment in 1979	−0.047	−0.018	0.018	4.840
	(0.013)	(0.012)	(0.026)	(0.089)
Usual hours in 1979	−1.98	−0.82	0.35	5.04
	(0.50)	(0.47)	(1.05)	(0.08)
Weeks of work in 1979	−2.47	−1.45	0.20	7.62
	(0.52)	(0.49)	(1.09)	(0.02)
Received public assistance in 1979	0.085	0.057	0.012	0.150
	(0.013)	(0.012)	(0.026)	(0.927)
Wage and salary income in 1979	−500.9	−393.3	−98.3	0.3
	(113.5)	(109.8)	(245.5)	(0.9)
Hours last week	−2.26	−1.45	−0.63	4.34
	(0.46)	(0.43)	(0.97)	(0.11)
Current employment	−0.064	−0.044	−0.019	8.740
	(0.012)	(0.011)	(0.025)	(0.013)
State fixed effects	No	Yes	Yes	
Demographic controls	No	Yes	Yes	
Instruments	No	No	3 QOB dummies	

Notes: For columns (1)–(3), the top number in each cell is the coefficient on the public school enrollment dummy in equation estimated using the indicated row variable on the left-hand side. Standard errors are reported in parentheses. Column (4) top number is χ^2 overidentification test statistic with associated p-value in parentheses. Sample size is 6,885.

programs, these women may be more willing to report income either to the tax authorities or to survey-takers.

In Table 8, I report estimated effects of public school enrollment on labor-market outcomes for female heads whose youngest child was younger than five. As in Table 7, the simple OLS results in column (1) are significant and wrong-signed. Adding the controls in column (2) drives the estimated coefficients toward zero and reduces their significance below conventional levels for rejecting a zero effect. By contrast to the results for single mothers whose youngest child is five, however, the column (3) IV estimates are uniformly insignificant, with four of the seven estimates wrong-signed. The IV estimates fail overidentification testing at the 0.01 level in two cases and the 0.10 level in two others, but they pass the overidentification test in two of the four wrong-signed cases. These results, together with the lack of a reduced-form relationship between outcomes and QOB, suggest that the most appropriate conclusion for single mothers whose youngest child is younger than five is simply that free schooling for the five-year-old has no effect on labor-market outcomes.

Tables 9 and 10 present estimates for married mothers. The OLS results generally suggest that labor supply of married women either does not respond or responds negatively to public school enrollment for the five-year-old. By contrast, the IV results again yield clear evidence of a significant positive enrollment effect on labor supply.[12] The levels of the IV coefficient estimates are generally smaller than those found for single mothers whose youngest child is five. However, baseline labor supply also is lower among the married mothers than among the single mothers, so the percentage effects are roughly comparable, ranging from 6 percent to 15 percent. With the exception of wage and salary income, it is not possible to reject equivalence of percentage effects for single mothers whose youngest child is five and either subpopulation of married mothers.[13]

Because married women on average have greater household income from sources other

[12] The IV specifications for married mothers pass overidentification tests at the 0.10 level in 13 of 14 cases.

[13] In the case of public assistance receipt, this conclusion is forced by the extremely large standard errors for the percentage response of married mothers, who are categorically ineligible for many welfare programs.

TABLE 9—EFFECT OF PUBLIC SCHOOL ENROLLMENT ON LABOR-SUPPLY MEASURES, MARRIED MOTHERS WHOSE YOUNGEST CHILD IS FIVE YEARS OLD

Left-Hand-Side Variable	(1) OLS	(2) OLS	(3) 2SLS	(4) Test of Overidentification
Employment in 1979	−0.001	0.009	0.048	2.56
	(0.004)	(0.004)	(0.008)	(0.28)
Usual hours in 1979	−0.26	0.15	1.71	2.10
	(0.16)	(0.16)	(0.30)	(0.35)
Weeks of work in 1979	−0.63	−0.18	1.71	1.18
	(0.20)	(0.19)	(0.37)	(0.55)
Received public assistance in 1979	0.007	0.003	0.002	1.56
	(0.001)	(0.001)	(0.002)	(0.46)
Wage and salary income in 1979	−283	−204	190	1.43
	(46)	(45)	(85)	(0.49)
Hours last week	−0.32	0.05	1.54	1.75
	(0.16)	(0.16)	(0.30)	(0.42)
Current employment	−0.000	0.009	0.050	1.19
	(0.004)	(0.004)	(0.008)	(0.55)
State fixed effects	No	Yes	Yes	
Demographic controls	No	Yes	Yes	
Instruments	No	No	3 QOB dummies	

Notes: For columns (1)–(3), the top number in each cell is the coefficient on the public school enrollment dummy in equation estimated using the indicated row variable on the left-hand side. White-robust standard errors are reported in parentheses. Column (4) top number is χ^2 overidentification test statistic with associated p-value in parentheses. Sample size is 53,163.

TABLE 10—EFFECT OF PUBLIC SCHOOL ENROLLMENT ON LABOR-SUPPLY MEASURES, MARRIED MOTHERS WHOSE YOUNGEST CHILD IS YOUNGER THAN FIVE YEARS OLD

Left-Hand-Side Variable	(1) OLS	(2) OLS	(3) 2SLS	(4) Test of Overidentification
Employment in 1979	0.004	0.014	0.031	3.72
	(0.004)	(0.004)	(0.008)	(0.15)
Usual hours in 1979	0.05	0.36	1.03	1.98
	(0.16)	(0.16)	(0.29)	(0.37)
Weeks of work in 1979	−0.13	0.19	1.47	3.34
	(0.17)	(0.17)	(0.32)	(0.19)
Received public assistance in 1979	0.011	0.007	0.000	0.18
	(0.001)	(0.001)	(0.003)	(0.91)
Wage and salary income in 1979	−152	−119	179	2.09
	(38)	(38)	(71)	(0.35)
Hours last week	0.06	0.29	1.21	0.72
	(0.15)	(0.14)	(0.27)	(0.70)
Current employment	0.003	0.010	0.037	0.53
	(0.004)	(0.004)	(0.008)	(0.77)
State fixed effects	No	Yes	Yes	
Demographic controls	No	Yes	Yes	
Instruments	No	No	3 QOB dummies	

Notes: For columns (1)–(3), the top number in each cell is the coefficient on the public school enrollment dummy in equation estimated using the indicated row variable on the left-hand side. White-robust standard errors are reported in parentheses. Column (4) top number is χ^2 overidentification test statistic with associated p-value in parentheses. Sample size is 52,134.

than their own labor supply, one would generally expect married mothers to react more elastically than single mothers to price subsidies. However, married mothers are not eligible for many forms of cash public assistance. Moreover, free public schooling provides a marginal price subsidy only for women whose optimal hours are no more than the length of the school day given that they work and enroll their five-year-old; for other women, at least part of the effect of enrollment comes in the form of an income subsidy. Differences across marital status in income elasticities of labor supply thus may help explain the similarity of the relative effects reported here.

III. Children Younger Than Five

Whether the maternal labor-supply response to child care costs differs with children's age has attracted little empirical attention, as authors generally have assumed a constant effect of child care prices. Two exceptions are Arleen Leibowitz et al. (1992) and Blau and Hagy (1998),[14] neither of which finds evidence that the effect of child care costs differs by children's age. Nonetheless, since most children who would be affected by extending preschool are younger than five, I consider briefly in this section the effects of school enrollment on three- and four-year-olds.

Among all mothers of three-year-olds in my sample, 23 percent report that their three-year-old is enrolled in school, with 7 percent enrolled in public programs. Among mothers of four-year-olds, the figures are 43 percent and 17 percent. In 1980, only a handful of state-funded public preschool programs existed.[15] An alternative source of enrollment is Head Start, which is generally administered at the site level and provides means-tested eligibility for children aged 3–5. In calculations not reported here, I compared age-specific Head Start enrollment statistics for FY1980 to cross tabulations of public school enrollment and family income relative to the federal poverty line. The results confirm that most public school enrollment among three- and four-year-olds in my sample is likely due to Head Start, particularly among single mothers.[16]

Because Head Start does not serve children younger than three, it is clear that QOB should be related to public school enrollment among three-year-olds. Even with such a low overall public school enrollment rate, I found a statistically significant first-stage relationship between enrollment and QOB among three-year-olds.

Since three-year-olds are eligible in principle, one might doubt the presence of a relationship between QOB and enrollment of four-year-olds. This doubt is unwarranted. Three-year-olds have typically made up only about one-fourth of Head Start enrollees (for FY1980 the figure was 24 percent). Because the program has never been fully funded, site-level Head Start directors have substantial discretion to give priority to older children. I have been unable to find any statistics on local age-enrollment policies, but a conversation with an official at the Head Start Bureau suggested that it is common for local administrators to restrict access to three-year-olds in order to serve more four-year-olds. As a result, a substantial fraction of younger four-year-olds likely were not able to enroll in Head Start based on their calendar date of birth.

The presence of de facto restrictions on four-year-old eligibility concerns only the first stage and so is empirically testable. In fact, I found that public school enrollment among four-year-olds follows the same monotonic pattern across QOB found among five-year-olds. Moreover, F-statistics testing the equivalence of public school enrollment across four-year-olds' QOB were 99 for single mothers and 390 for married mothers, so the relationship is highly significant whether or not the woman also has a child younger than four.

Given these facts, it is tempting to use QOB dummies as instruments for public school enrollment among three- and four-year-olds. However, enrollment of three- and four-year-olds in *private* school also increases with QOB. A likely explanation is that pricing and

[14] A third paper, Patricia M. Anderson and Philip B. Levine (1999), allows the effect of child care costs to vary with youngest child's age. However, this paper constrains the price coefficient to be the same over all preschool ages.

[15] According to Fern Marx and Michelle Seligson (1988), in 1980 only eight states either funded or had legislation related to pre-kindergarten programs.

[16] For example, 48 percent of all public school enrollment among four-year-old children of single mothers in my sample occurred for children living below 75 percent of the federal poverty line; virtually all of these children would have been income eligible for Head Start.

entry rules for private preschool programs reflect local public school entry rules, in order to accommodate eventual parental kindergarten enrollment decisions.[17] QOB will thus be related to marginal child care costs even for parents who do not enroll their children in publicly financed programs, so QOB dummies are inappropriate instruments for public school enrollment among younger children.

However, subject to the usual exclusion restrictions, QOB dummies will still be valid instrumental variables for *all* enrollment in either public or private programs. It is important to keep in mind that what is being identified is now a combination of the effects of free public school enrollment and reduced-price private enrollment due to aging of children.[18]

IV estimates yielded significant effects of all enrollment on employment in 1979, weeks of work in 1979, hours last week, and current employment of single mothers whose youngest child is four.[19] The percentage effects ranged between 20–60 percent, depending on the outcome. For single mothers with both a four-year-old and a younger child, there were significant effects for all but public assistance receipt. The percentage effects were quite large for these women, in several cases exceeding 100 percent; however, the confidence intervals on the larger percentage effects are very wide.[20]

Among single mothers whose youngest child is three, none of the estimates were significantly different from zero. Among single mothers with both a three-year-old and a younger child, significant estimates occurred for weeks in 1979, hours last week, and current employment. The percentage effects were over 100 percent for all three of these outcomes, but the confidence intervals (computed using the delta method for nonlinear functions of estimated parameters) on the percentage effects were all very large.

Among married mothers with both three- and four-year-olds (regardless of youngest child's age), enrollment has significant effects on all outcomes except public assistance receipt. The percentage effects range from 24 percent to 38 percent for three-year-old's enrollment among those whose youngest child is three, and from 18 percent to 27 percent for four-year-old's enrollment among those whose youngest child is four. For those with both a three-year-old and a younger child, all labor-supply estimates except for hours last week lie between 53 percent and 86 percent; the range is 25 percent to 52 percent for those with both a four-year-old and a younger child. For hours last week, the estimates were over 100 percent for married mothers who had either a three-year-old or a four-year-old and a younger child; however, the delta-method confidence intervals for both of these estimates were very large.

The results in this section show that access to public and cheaper private preschool programs generally leads to significantly greater labor supply for mothers of both three- and four-year-olds. While the estimates are not directly comparable to those for five-year-olds, they do make the case that large child care subsidies for parents of younger children are likely to have significant effects on maternal labor supply.

IV. Conclusion

The estimates reported in this paper suggest that public school enrollment among five-year-

[17] I thank Bill Evans for pointing this out to me.

[18] One might question my results for five-year-olds on the grounds just discussed. However, the decline in private enrollment rates of five-year-olds across QOB suggests that parents of five-year-olds use public school as a substitute for private school. Nonetheless, I reestimated the models reported in the previous section using all enrollment, rather than public school enrollment, as the endogenous right-hand-side variable. The estimates were uniformly about 25 percent larger for single mothers whose youngest child is five and continue to be insignificantly different from zero for single mothers who have both a five-year-old and a younger child. All enrollment estimates were about 40 percent larger than public school enrollment estimates among married mothers whose youngest child is five and about 25 percent larger among married mothers with both a five-year-old and a younger child.

[19] To conserve space, I do not report the estimates and percentage effects here; full tables are available on request.

[20] Because average labor supply among women with young children is very low, small absolute effects can imply very large relative effects. Also, because the percentage effect R is a nonlinear function of $\hat{\beta}$, its standard error rises with $\hat{\beta}$. Using the formula in footnote 10 and ignoring variance due to $\bar{y}_e$, the standard error for R is $[\bar{y}_e/(\bar{y}_e - \hat{\beta})^2]\sqrt{\hat{V}(\hat{\beta})}$. Holding constant $\bar{y}_e$, the width of the confidence interval for R is thus increasing in $\hat{\beta}$ itself. When the effect is large, estimates will thus be very noisy.

olds is endogenously related to labor-market outcomes. Instrumental variables results show that public school enrollment for five-year-olds has a statistically and economically significant effect on labor-market outcomes among single mothers whose youngest child is five years old as well as all married mothers with a five-year-old. The QOB instrumental variables are very powerful instruments and generally pass over-identification tests.

For mothers of three- and four-year-olds, the effects of public schooling cannot be separated from the effects of cheaper private programs. However, with the exception of single mothers whose youngest child is three, estimates of the combined effects generally suggest large and significant labor-supply effects of reduced child care costs.

In summary, these results suggest that large child care subsidies are likely to increase maternal labor supply among a significant fraction of women with preschool children.

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