Predicting Female Labor Supply

Effects of Children and Recent Work Experience

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ABSTRACT

This study examines the empirical associations between female labor supply and child status and marital status using 1970 and 1980 U.S. census data and 1971 and 1981 Canadian census data. When the data are used in a purely cross-sectional manner, without controlling for previous labor supply, we find, as others have, that female labor supply is negatively related to the number of children a woman has had. However, this relationship changes when we condition on weeks of work in the previous year.

This study makes use of longitudinal information in the Canadian and U.S. census data that has been largely ignored. The paper also explores certain econometric issues raised by the nature of the empirical results.

I. Introduction

Child status variables account for the largest share of the explained variation in most empirical models of female labor supply. Children have also been a focus in theoretical discussions of female labor supply, and in proposed explanations for why the work behavior and wages of women and men differ.

According to the dominant economic model, the labor supply of individuals is

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determined by the intersections of their reservation and offered wage functions: an individual's reservation wage being the compensation required for the individual to be willing to work one (or one more) unit time period, such as an hour, and the offered wage being what an employer would be willing to pay for this labor input. With this theoretical context, factors that act to raise the reservation wage or lower the offered wage of an individual will tend to decrease his or her labor supply. In largely separate streams of the literature on female labor supply and earnings, economists have argued that family responsibilities, and particularly children, will affect both the reservation and offered wage rates of women and that the impacts on their labor supply will be predominantly negative.[1]

In Section II, we describe the census data sets used in this study for the United States and Canada. We also show cross-sectional descriptive evidence of the sort that has been interpreted as confirmation of the traditional views of economists concerning how children affect female labor supply. In Section III we examine patterns in the reference week employment rates for women (the employment rates for the week prior to enumeration) conditional on weeks of work in the previous calendar year. The empirical results motivate the presentation in Section IV of reference week employment rates that are also conditional on parity and the presence of a child under six years of age. The results in Section IV are not entirely consistent with traditional views of how children affect the work behavior of women, and appear to support the position that child status variables that are directly included in models of female labor supply will serve as proxies for omitted factors in addition to capturing the direct effects of the time and other resources required to bear and raise children.

Section V explores certain econometric issues related to the empirical findings. If child status variables that are directly included in labor supply equations serve as proxies for omitted factors, the estimated coefficients for the child status variables will be biased estimates of the impacts of exogenous changes in child status on female labor supply. This *coefficient bias* problem could lead to wrong inferences concerning the impact on female labor supply of government programs (or other exogenous circumstances) affecting fertility behavior. At the same time, however, the *prediction bias* for a labor supply equation will be reduced if included child status and other explanatory variables are able to account, at least partially, for the effects of important and persistent unobservable factors affecting individual labor supply behavior. Many of the suggested reasons for interest in the work behavior of women imply a need for accurate predictions of the labor supply of various types of women. Also, biased predictions that happen to reinforce erroneous popular perceptions about the labor supply behavior of women with children can contribute to, or serve as a justification for, statistical discrimination against women in the workplace. The tradeoffs between coefficient and prediction biases, for both directly included and instrumental child status variables, are explored in the context of a simplified "true" model of female labor

1. See Nakamura and Nakamura (1992); Browning (1992); Siegers, de Jong-Gierveld and van Imhoff (1991); and Lehrer and Nerlove (1984, 1986) for surveys of literature on the effects of children on female labor supply.

supply that is consistent with the empirical findings presented in the earlier portions of the paper.

Section VI concludes.

II. Marriage, Children, and Female Labor Supply: Cross-Sectional Evidence

The hypothesized negative effects of marriage and children on female labor supply are evident in simple cross-tabulations such as those presented in the following subsections for two alternative measures of labor supply: weeks of work in a year, and employment status (employed or not employed) in a week. It is important to note that we do not distinguish women who are employed and at work from those who are employed and not at work. Rather, we use the term "worked" as meaning employed. This study provides no information on the changes examined by Klerman and Leibowitz (1993, 1994) in the propensity to be at work for women who are employed—an important dimension of female labor supply in the months immediately prior to and following childbirth.

Our data sources for this study are public use sample files from the 1970 and 1980 U.S. censuses of population, and from the 1971 and 1981 Canadian censuses. Our computations are based on the observations for women 25–45 years of age who graduated from high school but did not attend college or university and for whom information is available for weeks of work in the previous calendar year, employment status in the reference week, marital status, age, number of children ever born, and age (or age group) of the youngest child for the married women with children. Our U.S. data samples were further restricted to white women. (See the Data Appendix for further details.) By a combination of our data selection criteria and cross-classifications, we control for the main variables found to be empirically important in previous studies of female labor supply.

A. Marriage and Child Status Effects on Weeks of Work

Table 1 shows distributional information for women by weeks of work in the calendar year preceding each of the designated censuses. The specified weeks of work categories are 0, 1–26, 27–47, and 48+ weeks, for comparability with other findings presented in Nakamura and Nakamura (1992) for U.S. women.[2]

In both the top and bottom panels of Table 1, the figures are arranged in three groups of three rows each: a first row giving 1970 or 1971 figures, a second giving 1980 or 1981 figures, and a third giving the 1980/81–1970/71 differences. Each three rows are for one of our demographic groupings of U.S. or Canadian women: (1) women who at the time of enumeration were never married, divorced or widowed and who had no children ever born if divorced or widowed; (2) married women with no children ever born; and (3) married women with children ever born. These three demographic groups will be referred to, respectively, as unmar-

2. See Section 3 of the Data Appendix for details of the weeks categories for the 1971 Canadian census data.

Table 1
Distribution for Annual Weeks of Work

Marital status	Child status	Census year	Weeks of work in previous year 0	1–26	27–47	48+
		U.S. Women				
Not married	No children	1970	.06	.09	.14	.70
		1980	.07	.10	.14	.68
		80–70	.01	.01	.00	−.02
Married	No children	1970	.19	.12	.16	.53
		1980	.14	.12	.14	.60
		80–70	−.05	.00	−.02	.07
Married	Children	1970	.56	.15	.09	.20
		1980	.39	.16	.13	.32
		80–70	−.17	.01	.04	.12
		Canadian women				
Not married	No children	1971	.10	.14	.21	.54
		1981	.08	.17	.17	.58
		81–71	−.02	.03	−.04	.04
Married	No children	1971	.12	.12	.21	.54
		1981	.09	.12	.15	.64
		81–71	−.03	.00	−.06	.10
Married	Children	1971	.53	.16	.13	.17
		1981	.36	.18	.13	.33
		81–71	−.17	.02	.00	.16

ried and childless, married and childless, and married with children. The presumption has been that family needs for a woman's time tend to increase in moving from group 1 to group 2 to group 3.

In Table 1, the marital and child status related differences in labor supply can be seen most clearly by focusing on the 0 and 48+ weeks columns: Columns 1 and 4.

For U.S. women, moving from those who are unmarried and childless to those who are married and childless to those who are married with children, labor supply measured in terms of weeks of work falls. In particular, the percentage of women with 0 weeks of work rises steeply (6 to 19 to 56 percent for the 1970 census data; 7 to 14 to 39 percent for the 1980 data) and the percentage who worked 48+ weeks falls (70 to 53 to 20 percent for the 1970 data; 68 to 60 to 32 percent for the 1980 data).

Comparing the 1980 with the 1970 U.S. census figures, there is little evidence of change for the unmarried, childless women. However, for the married, childless women and the married women with children, labor supply increased. The proportions of married women who worked 48+ weeks rose and the proportions who worked 0 weeks fell. This is in accord with the labor supply findings reported by others, and is one reason why married women have been the focus of female labor supply research. The increases evident in Table 1 in the proportions of married women who worked 48+ weeks and the decreases in the proportions who worked 0 weeks are particularly large for the married women with children.

The figures for Canadian women in the bottom panel of Table 1 exhibit similar patterns to those noted for the United States.

For the Canadian women, labor supply decreases in moving from those who are unmarried and childless to those who are married and childless to those who are married with children. In particular, the percentage with 0 weeks of work rises (10 to 12 to 53 percent for the 1971 data; 8 to 9 to 36 percent for the 1981 data). Also, the percentage who worked 48+ weeks is lower for the married women with children compared with the childless married women (17 percent for those with children compared with 54 percent for those without for the 1971 data; 33 percent for those with and 64 percent for those without for the 1981 data). In contrast to the U.S. results, however, the percentage figures for the unmarried, childless women are not higher than for the married childless women.

Comparing the 1981 with the 1971 Canadian census figures, the main changes are increases in the proportions who worked 48+ weeks for married, childless women, and particularly for married women with children. Also there are decreases in the 0 weeks proportions. These decreases are most pronounced for married women with children. Thus the changes over time in the Canadian figures are qualitatively similar to the changes discussed above for the United States.

The main conclusions emerging from Table 1 can be summarized as follows:

(1) In both the United States and Canada, women seem to supply less labor as family responsibilities increase. In particular, using census data for 1970 and 1980 for the United States and for 1971 and 1981 for Canada, we find that in moving from unmarried, childless women to married, childless women to married women with children, the proportion of women who worked 0 weeks rises and the proportion who worked 48+ weeks usually falls.

(2) There were no numerically substantial changes over the 1970/71 to 1980/81 period in the cross-sectional weeks of work distributions for unmarried, childless women. In both countries, however, there were increases for married women in the proportions who worked 48+ weeks and decreases in the proportions with 0 weeks of work. These changes were particularly large for the married women with children.

B. Marriage and Child Status Effects on Reference Week Employment Rates

Table 2 gives the reference week employment rates for the same women whose weeks of work distributions are shown in Table 1. In Table 2, the women have been grouped by age (20–24, 25–29, 30–45) as well as by marital and child status. The layout of Table 2 is similar to Table 1 except that now the columns are for the three age subgroups.

For each age subgroup for both the United States and Canada, and for the 1970/71 as well as the 1980/81 census years, the reference week employment

Table 2
Reference Week Employment Rates

Marital status	Child status	Census year	Age of women 20–24	25–29	30–45
		U.S. women			
Not married	No children	1970	.85	.91	.85
		1980	.84	.87	.85
		80–70	−.01	−.04	.00
Married	No children	1970	.69	.64	.57
		1980	.80	.76	.69
		80–70	.11	.12	.12
Married	Children	1970	.23	.27	.39
		1980	.38	.41	.55
		80–70	.15	.14	.16
		Canadian women			
Not married	No children	1971	.79	.83	.88
		1981	.82	.89	.87
		81–71	.03	.06	−.01
Married	No children	1971	.76	.76	.65
		1981	.82	.84	.89
		81–71	.06	.08	.15
Married	Children	1971	.31	.29	.38
		1981	.41	.47	.57
		81–71	.10	.18	.19

rates decline moving from unmarried, childless women to married, childless women to married women with children. Thus, for this measure too, it seems that labor supply declines as family responsibilities increase.

The age related patterns evident in Table 2 also lend credibility to conceptualizations of female labor supply as reflecting period by period tradeoffs between work time and family time. Moving from the 20–24 to the 25–29 to the 30–45 age group, notice that for the United States for both 1970 and 1980 and for Canada for both 1971 and 1981, the reference week employment rates for married women with children rise. This could reflect declines in child care responsibilities as women age through their 20s to their 30s and 40s.

As in Table 1, in Table 2 we find that the cross-sectional labor supply behavior of unmarried, childless women has been relatively stable over the 1970/71 through 1980/81 period. We find too that there have been substantial increases in the reference week employment rates for married women—particularly for married women with children.

III. Conditioning on Prior Work Status

A. Current versus Previous Period Labor Supply Choices

Every woman either did or did not work in the immediate past. This is the context for the real life choices women make about current period work behavior. Those women who have been working may have jobs they could continue in without incurring search costs, and will already have found ways of dealing with transportation, child care needs, family and personal feelings, and other complications arising from their jobs. Women who have not been working would usually have to search for jobs in order to start working, and would have to solve the practical complications working women face. These are some of the reasons we would expect to find that women who were working in the immediate past will be far more likely to be employed in the present than those who had not been working.

It is possible to use census data to examine the current period labor supply of women conditional on labor supply in the previous year.[3] The 1970 and 1980 public use samples for the United States and the 1971 and 1981 public use samples for Canada provide information on both employment status in the reference week (the week preceding the census enumeration week) and on weeks of work in the preceding calendar year.

If being married and having children reflect greater current period family demands for a woman's time, we might expect to find that among the women who have not been working (zero weeks of work in the previous year), those with greater family responsibilities will be less likely to start working and hence will have lower reference week employment rates. Likewise, we might expect that among the women who had been working (1–26, 27–47, or 48+ weeks of work in the previous year), those in a given previous weeks of work group with greater family responsibilities will be more likely to quit their jobs, and hence will have lower reference week employment rates than those with lesser family responsibilities. Relevant empirical evidence is examined in the following subsection.

B. Reference Week Employment Rates for Women Classified by Weeks of Work in the Preceding Year

Reference week employment rates computed using our U.S. and Canadian data samples are shown in Table 3 for women classified by weeks of work in the previous year (0, 1–26, 27–47, 48+) as well as by marital and child status and by age. The layout differs from Tables 1 and 2. The three columns of employment rate figures in the left half of the table are for the U.S. women. The first column is for the unmarried, childless women. The second column is for the married,

3. Nakamura and Nakamura (1992) show empirical results based on this same approach for 1980 U.S. census data. Results are presented in that study for women with less than high school educations and for those with post-secondary schooling as well as for high school graduates, and similar patterns are found for all three schooling groups.

Duleep and Sanders (1994) use this same approach with 1980 U.S. census data for immigrant women, categorized by country of origin, and for native born women. The patterns reported are strikingly similar for all of the immigrant groups and the native born women.

childless women. And the third is for the married women with children. Likewise the three columns in the right half of the table are for our three marital-child status groups of Canadian women, ordered from left to right in terms of increasing family responsibility.

In Table 3, there is a separate panel for each classification of weeks of work in the preceding year: 0, 1–26, 27–47, 48+. Within each panel, the three groups of three rows each are for our age subgroups: 20–24, 25–29, and 30–45, respectively.

For each age-previous weeks of work grouping in Table 3 for each of the U.S. and the Canadian census years, we find that the reference week employment rates usually decline as expected in moving from the unmarried, childless women to the married, childless women to the married women with children. For example, from row 2 of panel 2 for women who worked 1–26 weeks in the previous year, we see that for 1980 for U.S. women 20–24, the reference week employment rate is .61 for those who were unmarried and childless, .57 for those who were married and childless, and .38 for those who were married with children; and for 1981 for Canadian women 20–24, the reference week employment rate is .68 for those who were unmarried and childless, .61 for those who were married and childless, and .38 for those who were married with children.

The age patterns evident in Table 3 for married women with children are also consistent with traditional views on the role of child status variables in models of female labor supply. The reference week employment rates for those with positive weeks of work in the previous year tend to increase moving from the 20–24 to the 25–29 to the 30–45 age group. This pattern is what might be expected of women as their childbearing years come to an end and their children begin to grow up.

The differences in the levels of the reference week employment rates for women grouped by weeks of work in the previous year are large compared with the marital and child status related differences within each of the prior weeks of work groupings. One possible reason for this that would bolster traditional views of the information conveyed by marital and child status attributes is that, on average, those married women with children who worked greater numbers of weeks in the previous year have fewer, and perhaps older, children than those who worked fewer weeks. This possibility is explored in Section IV by grouping married women by parity, and also by separately examining the employment behavior of those married women with a child younger than six.

Note, however, that for married women with children who did not work in the previous year (that is, for those with zero weeks in the previous year), the reference week employment rates *decrease* rather than increase with age except for Canada for 1971.

Note also that, after grouping by weeks of work in the previous year, it is *not* primarily the married women for whom we observe substantial changes in reference week employment rates from 1970/71 to 1980/81. After controlling for weeks of work in the previous year, the evidence in Table 3 of increases over time in employment rates for married women with children is quite modest—particularly for the United States. On the other hand, for both the United States and Canada, there were quite large increases over this time period in the employment rates for unmarried, childless women who did not work in the previous year. This

Table 3
Reference Week Employment Rates Conditional on Weeks of Work in Previous Year

Age group	U.S. census	U.S. women: No children ever born, Not Married	U.S. women: No children ever born, Married	U.S. women: Children, Married	Canadian census	Canadian women: No children ever born, Not Married	Canadian women: No children ever born, Married	Canadian women: Children, Married
			Worked 0 weeks in previous year					
20–24	1970	.13	.17	.07	1971	.30	.12	.04
	1980	.23	.22	.10	1981	.36	.41	.12
	80–70	.10	.05	.03	81–71	.06	.29	.08
25–29	1970	.14	.11	.04	1971	.23	.20	.05
	1980	.16	.14	.07	1981	.37	.25	.11
	80–70	.02	.03	.03	81–71	.14	.05	.06
30–45	1970	.00	.05	.04	1971	.19	.08	.06
	1980	.13	.05	.06	1981	.37	.15	.11
	80–70	.13	.00	.02	81–71	.18	.07	.05
			Worked 1–26 weeks in previous year					
20–24	1970	.66	.46	.38	1971	.67	.55	.37
	1980	.61	.57	.38	1981	.68	.61	.38
	80–70	−.05	.11	.00	81–71	.01	.06	.01

25–29	1970	.73	.43	.41	1971	.53	.53	.41
	1980	.58	.52	.40	1981	.75	.62	.47
	80–70	−.15	.09	−.01	81–71	.22	.09	.06
30–45	1970	.60	.51	.54	1971	.79	.44	.49
	1980	.58	.60	.56	1981	.72	.59	.60
	80–70	−.02	.09	.02	81–71	−.07	.15	.11
Worked 27–47 weeks in previous year								
20–24	1970	.79	.60	.54	1971	.84	.74	.57
	1980	.79	.69	.57	1981	.86	.78	.62
	80–70	.00	.09	.03	81–71	.02	.04	.05
25–29	1970	.88	.62	.72	1971	.89	.74	.60
	1980	.85	.79	.70	1981	.83	.81	.71
	80–70	−.03	.17	−.02	81–71	−.06	.07	.11
30–45	1970	.80	.75	.83	1971	.78	.80	.83
	1980	.85	.81	.80	1981	.86	.85	.80
	80–70	.05	.06	−.03	81–71	.08	.05	−.03
Worked 48+ weeks in previous year								
20–24	1970	.96	.85	.77	1971	.94	.90	.75
	1980	.95	.96	.83	1981	.95	.94	.76
	80–70	−.01	.11	.06	81–71	.01	.04	.01
25–29	1970	.98	.90	.91	1971	.91	.92	.79
	1980	.97	.93	.93	1981	.97	.94	.86
	80–70	−.01	.03	.02	81–71	.06	.02	.07
30–45	1970	.98	.97	.95	1971	.97	.96	.93
	1980	.98	.96	.95	1981	.96	.97	.93
	80–70	.00	−.01	.00	81–71	−.01	.01	.00

suggests that in trying to understand the increases over time in the labor supply of married women with children that are evident in cross-sectional data of the sort presented in Tables 1 and 2, more attention should be focused on the year-to-year patterns of labor supply prior to marriage and prior to the start of child bearing.[4]

IV. Conditioning on Prior Work Status and Child Status

A. Parity Specific Reference Week Employment Rates

In Table 4 we show parity specific reference week employment rates for the married women in our estimation data samples grouped by weeks of work in the previous year. The four columns on the left for the United States and the corresponding four columns on the right for Canada are for married women with 0, 1, 2, and 3+ children ever born, respectively.

In all cases, the reference week employment rates for the married women with one child (Columns 2 and 6) are less than the corresponding values for the married women with no children (Columns 1 and 5, respectively). This is consistent with the traditional view that children have negative effects on the amount of labor their mothers supply.

However, looking at Columns 2–4 and 6–8, we find that for married women with 1, 2, and 3+ children, the reference week employment rates *rise* with the number of children, except for the women who did not work at all in the previous year. The increases are particularly large for the women who worked 1–26 or 27–47 weeks in the previous year. For example, for the United States, the reference week employment rates for married women who worked 27–47 weeks in the previous year are .58 in 1970 and .66 in 1980 for those with one child ever born compared with .80 in both 1970 and 1980 for those with 3+ children! Similarly, for Canada, the employment rates for married women who worked 27–47 weeks in the previous year are .58 in 1971 and .69 in 1981 for those with one child compared with .81 in 1971 and .80 in 1981 for those with 3+ children.

One might speculate that the observed increases with parity in the reference week employment rates are a child maturation or family completion effect. The traditional view has been that many women might wait until they felt their families were complete and their youngest children were growing up before returning to, or devoting more time again to, market work, and higher proportions of the women in our estimation samples with 2 or 3+ children may have completed their families. Besides, a woman with 3+ children whose youngest child is, say, 18 may well have less family responsibilities than a woman with a single child still in the preschool years. However, the information in the following subsection does not support these explanations of the parity related increases in labor supply evident in Table 4.

4. Shaw (1994) also presents evidence supporting this perspective. See Shapiro and Mott (1994) as well.

Table 4
Probability of Work in the Reference Week Conditional on Weeks of Work in Previous Year for Married Women Grouped by Number of Children Ever Born

	United States				Canada			
	Number of children ever born							
Census	0	1	2	3+	0	1	2	3+
	Worked 0 weeks in previous year							
1970/71	.08	.06	.04	.04	.12	.06	.05	.05
1980/81	.12	.08	.06	.06	.27	.14	.10	.10
	Worked 1–26 weeks in previous year							
1970/71	.47	.33	.46	.52	.53	.38	.42	.51
1980/81	.56	.42	.50	.51	.60	.48	.55	.55
	Worked 27–47 weeks in previous year							
1970/71	.63	.58	.77	.80	.75	.58	.80	.81
1980/81	.74	.66	.76	.80	.81	.69	.78	.80
	Worked 48+ weeks in previous year							
1970/71	.89	.88	.94	.93	.91	.83	.90	.91
1980/81	.95	.91	.95	.94	.94	.87	.91	.91

B. Parity Specific Reference Week Employment Rates for Women with a Child Younger than Six

For Table 5, we have separated out the married women with at least one child younger than six years old. For this subsample of women, reference week employment rates for those with 1, 2, or 3+ children ever born are shown for the United States (Columns 1–3) and Canada (Columns 4–6).

We still find that the reference week employment rates generally *rise* with the number of children, except for the women who did not work at all in the previous year. And, again, we find that the parity related increases in the reference week employment rates are particularly large for the women who worked 1–26 or 27–47 weeks in the previous year.

What we find from both Tables 4 and 5 for married women who worked in the previous year is that those with three or more children tend to be *more* committed to working than those with two children, and likewise for those with two children compared with those with one child. For women who worked in the previous year, the employment rates for the 3+ parity women—even those with a child under six—are sometimes higher than for the married women with no children ever born. The tendency of reference week employment rates to rise with parity, after controlling for weeks of work in the previous year, has been demonstrated

Table 5
Probability of Work in the Reference Week Conditional on Weeks of Work in Previous Year for Married Women With a Child Younger Than Six Grouped by Number of Children Ever Born

	United States			Canada		
	Number of children ever born					
Census	1	2	3+	1	2	3+
	Worked 0 weeks in previous year					
1970/71	.06	.04	.03	.06	.03	.05
1980/81	.08	.06	.05	.13	.09	.08
	Worked 1–26 weeks in previous year					
1970/81	.27	.37	.49	.34	.37	.43
1980/81	.35	.46	.39	.44	.47	.51
	Worked 27–47 weeks in previous year					
1970/71	.52	.70	.72	.50	.59	.58
1980/81	.58	.70	.71	.66	.73	.77
	Worked 48+ weeks in previous year					
1970/71	.78	.86	.95	.71	.74	.83
1980/81	.88	.94	.91	.83	.87	.83

using 1980 U.S. census data even for married women with a youngest child under one year of age (Nakamura and Nakamura 1992, Tables 13, 14, 15, pp. 59–61). Unfortunately, the Canadian data do not allow further breakdown of women with a youngest child under six by the age of the child.

Perhaps Paul Schultz has been right in arguing that, in addition to capturing the family time demands on women that come with having children, the child status variables in family labor supply models also capture tastes for a home oriented versus a career oriented lifestyle and other preexisting conditions that were *not caused* by having children to care for.[5]

The presence of a preschool child does seem to reduce the period to period conditional labor supply of women, as would be expected. Comparing the Table 5 figures with the corresponding figures in Table 4 for women with 1, 2, or 3+ children, the Table 5 figures for the married women with a child younger than six do tend to be lower, particularly for the women who worked 27–47 or 48+ weeks in the previous year. For example, the Table 5 reference week employment rate figures for 1, 2, and 3+ parity U.S. women who worked 48+ weeks in the previous year are .88, .94, and .93, respectively, for 1970 and .91, .95 and .94 for

5. See, for example, Schultz (1974).

1980, while the corresponding Table 5 figures for married women with at least one child younger than six are .78, .86, and .95 for 1970 and .88, .94, and .91 for 1980.

The evidence in Tables 4 and 5 should encourage more research (and greater caution in making assumptions and assertions) concerning how children affect female labor supply. We also need a better understanding of the other factors that child status variables pick up, in a proxy sense, when they are directly included in models of female labor supply behavior.

Paul Schultz has made no use of information about previous work behavior in his research on female labor supply. Nevertheless, the evidence in Tables 4 and 5 suggests extending Schultz's arguments that child status variables may serve as proxies for tastes and other unobservable preconditions to allow for the possibility that the nature and magnitudes of these proxy effects may differ systematically for women who worked different numbers of weeks in the previous year. Our evidence suggests that the child status and previous weeks of work attributes of women may serve as joint indicators of persistent unobservable factors affecting current labor supply. For example, it may be that the women with three or more children who also worked 48+ weeks in the previous year tend to be women with particulrly strong career desires, or pressing and persistent unobserved needs for additional income. The material in the following section builds on this insight.

V. Econometric Implications

In Section IV we found that when married women are grouped by parity, for those who worked in the previous year the current employment rates tend to rise rather than fall with increases in the number of children ever born. This is true even when the analysis is limited to those women with a child younger than six at home.

Among other things, these findings suggest that the child status variables are picking up the effects of tastes and other preconditions that affect a woman's labor supply from one period to the next; not just the direct effects of children on a family's needs for a woman's time. If child status variables are serving as proxies for omitted factors, and if labor economists wish to interpret the child status coefficients in female labor supply equations as reflecting the effects on current labor supply of exogenous changes in child status, then direct inclusion of child status variables in models of female labor supply will lead to biased estimates of the child status response coefficients. This coefficient bias problem is formally demonstrated in subsection B below in the context of a simplified "true" model defined in subsection A. Related prediction bias problems that can contribute to statistical discrimination against working women with children are discussed in subsections C and D.

A commonly proposed solution to the problem of biased estimates of the coefficients of directly included child status variables is to use instrumental child status variables. Finding appropriate instruments is difficult, as Lehrer (1992) explains. But even if perfect instruments were available, this approach could worsen rather than ameliorate the prediction problems that can contribute to statistical discrimination against working women with children. This is demonstrated in subsection E.

Subsection F discusses an alternative proxy approach.

A. A Simplified "True" Model

Suppose that there are unobservable factors that affect women's labor supply behavior period after period and may also have affected their fertility choices, such as long standing tastes for a career versus a home oriented lifestyle. Or suppose there are unobservable factors tied to previous employment status such as seniority related wage benefits and the option to work without expending resources on job search. We will denote unobservable tastes and other preconditions affecting current labor supply that are either persistent or are associated with work behavior in the previous period by T. A woman's current labor supply (measured perhaps by annual weeks or the probability of being employed in a week) will be denoted by h. C is used to denote child status. For simplicity, C can be thought of in this section as the number of children ever born. X is a measure of all other observable factors affecting a woman's labor supply.

Suppose that a woman's labor supply is determined by the following "true" data generating equation:

$$(1) \quad h = \alpha_0 + \alpha_1 C + \alpha_2 X + \alpha_3 T + \varepsilon,$$

where ε is an error term with mean zero that is distributed independently of C, X, and T. Suppose also that ε is uncorrelated over time for each woman. The slope coefficients α_1, α_2, and α_3 are specified to be true behavioral response parameters. For instance, α_1 is specified to represent the expected change in h directly attributable to a one unit change in C.

Suppose we know that there are interrelationships among the unobservable variable T and the observable factors C and X in (1). For instance, suppose that women who have wanted a career tend to have higher values of T year after year and lower values of C. Impacts of T on labor supply in previous periods will, of course, be embedded in lagged labor supply variables. Thus, if the values of T are persistent, lagged labor supply variables might serve as proxies for T in an equation for current labor supply. Also, women who worked in the previous year may tend to have higher values of T because of the precondition that they have a job they can continue in without incurring search costs. For expositional simplicity, in this section, we will categorize *work in the previous year* using a single dummy variable W, where W equals 1 if a woman had positive weeks of work in the previous year and equals 0 otherwise.

The population of values for any variable can always be represented as the predicted values for the population ordinary least squares (OLS) regression of that variable on *any* other specified set of variables plus the corresponding population of OLS residual values. By construction, the OLS residuals will have a mean of zero and will be orthogonal to the explanatory variables included in that particular population regression model. Thus, we can represent the hypothetical population of values for the unobservable tastes and preconditions variable by

$$(2)\quad T = \beta_0 + \beta_W W + \beta_C C + \sum_{I=1}^{3} \beta_{WCI} WCI + \beta_X X + v,$$

where the βs are population regression parameters and v is the OLS residual term. Equation (2) is simply one way of summarizing the population intercorrelations between the tastes and preconditions variable, T, and the variables on the right hand side of (2). The variables WCI for $I = 1, 2, 3$ on the right hand side of (2) are the product of the lagged work status dummy variable, W, and child status dummy variables, with $C1 = 1$ if a woman has one child (hence $C = 1$), $C2 = 1$ if she has two children (hence $C = 2$) and $C3 = 1$ is she has three or more children (hence $C \geq 3$).

In (2), the OLS residual term v will be uncorrelated by construction with the lagged work status variable W, the child status variable C, the product variables WCI for $I = 1, 2, 3$, and the observable factors variable X. This aspect of (2) does not depend on any assumed properties of our simplified "true" model. The added property that *is* assumed as part of our simplified scenario is that v is random over time for each woman; that is, in our simplified true world, we assume that the variables W, C, WCI for $I = 1, 2, 3$, and X fully account (in at least a proxy sense) for the persistence over time for individual women of unobservable tastes and preconditions affecting labor supply.

The model given by (1) and (2) can accommodate the behavioral patterns exhibited in Table 4 if $\beta_W > 0$, $\beta_C < 0$, $0 < \beta_{WC1} < |\beta_C|$, $\beta_{WC2} > 2\,|\beta_C|$, and $\beta_{WC3} > 3\,|\beta_C|$ in (2), and if $\alpha_1 < 0$ and $\alpha_3 > 0$ in (1). However, the relationships developed below do not depend on whether the parameters of (1) and (2) satisfy these qualitative properties.

Substituting (2) into (1) yields the true reduced form model for h in terms of the observable variables in (1) and (2):

$$(3)\quad h = (\alpha_0 + \alpha_3\beta_0) + \alpha_3\beta_W W + (\alpha_1 + \alpha_3\beta_C)C + \sum_{I=1}^{3} \alpha_3\beta_{WCI} WCI + (\alpha_2 + \alpha_3\beta_X)X + [\alpha_3 v + \varepsilon].$$

In our simplified world, the error terms ε, v, and $[\alpha_3 v + \varepsilon]$ are uncorrelated with the right hand variables in (3) and are also random over time for individual women.

B. Coefficient Bias Problems Due to Omitting T

In the context of our simplified world, if T were observable, then unbiased (or consistent) estimates of the labor supply response parameters α_1 and α_2 in (1) could be obtained simply by regressing h on C, X, and T. However, T is not observable.

If h is regressed on C and X ignoring T, the estimated coefficients will be unbiased estimates of the coefficients (given in parentheses) of

$$(4)\quad h = (\alpha_0 + \alpha_3 b_{T\cdot C,X}) + (\alpha_1 + \alpha_3 b_{T,C\cdot X})C + (\alpha_2 + \alpha_3 b_{T,X\cdot C})X + [\alpha_3 e_{T\cdot C,X} + \varepsilon].$$

The "b components" of the coefficients in (4) have a regression interpretation. T, C, and X are jointly distributed. As already noted, the population of values for any one of these variables can be represented as the population OLS regression of the designated variable on the others plus the corresponding residual values for the population regression. Hence we can represent T as

$$(5) \quad T = \{b_{T \cdot C,X} + b_{T,C \cdot X}C + b_{T,X \cdot C}X\} + e_{T \cdot C,X},$$

where the b's are the population OLS coefficients and $e_{T \cdot C,X}$ is the OLS residual term which will be uncorrelated by construction with C and X. However, unlike v in (2), there is no methodological reason or any basis in the assumed properties of our simplified model for claiming that $e_{T \cdot C,X}$ will be random over time for individual women. On the contrary, if β_W, β_{WC1}, β_{WC2}, and β_{WC3} in (2) are nonzero, then we would expect the values of $e_{T \cdot C,X}$ to be persistently high for some women and low for others.

The subscripts used in (5), and in a number of subsequent expressions, have the dependent variable first (T for this equation). When the dependent variable is followed by a comma and then another variable, the coefficient is for that other variable. Additional variables that are controlled for in the regression model follow a dot. Hence $b_{T,C \cdot X}$ is the coefficient of C in an equation for T which also includes X. Note that Equation (4) can be obtained by substituting (5) into (1), just as (3) can be obtained by substituting (2) into (1).[6]

For purposes of discussion, we will assume that the coefficient of the child status variable in (5), $b_{T,C \cdot X}$, is negative, just as we assumed that β_C in (2) is negative. This is in accord with empirical findings of a negative relationship between labor supply and the number of children when a continuous variable for number of children is entered as a linear term with no account taken of previous work behavior.

Suppose that the purpose in regressing h on C and X is to obtain estimates of the true behavioral response parameters α_1 and α_2 in (1). From this perspective, the terms

$$(6) \quad \alpha_3 b_{T,C \cdot X} \text{ and } \alpha_3 b_{T,X \cdot C}$$

in the coefficients of (4) are *coefficient biases:* the type of biases that Paul Schultz has argued are a problem when child status variables are directly entered into models of female labor supply. If α_3, the coefficient of T in (1), is positive and $b_{T,C \cdot X}$ is negative, then we see that an estimated version of (4) will tend to over estimate the direct negative impact of having children to care for on female labor supply (that is, $\alpha_3 b_{T,C \cdot X}$ will be negative).

C. Prediction Bias Problems

A different sort of bias problem are the biases in predicting labor supply for women with different child status attributes and who did or did not work in the

6. This interpretation of bias components due to correlations between the error term and the explanatory variables can also be arrived at more conventionally. Consider the model $y = \beta X + u$. The OLS coefficient vector for the regression of y on X is $\hat{\beta} = (X'X)^{-1}X'y$, and $E(\hat{\beta}) = \beta + E[(X'X)^{-1}X'u]$. But $E[(X'X)^{-1}X'u]$ is just the OLS coefficient vector for the population regression of u on X.

previous year. This prediction problem is of interest, for example, to employers considering hiring or training or promoting women who are of child bearing ages.

The *prediction bias* problem for Equation (4) can be analyzed by comparing (4) with the true reduced form Equation (3). Subtracting the systematic portion of (3) from the systematic portion of (4), we see that the *prediction bias* when h is directly regressed on C and X is

$$(7)\quad \alpha_3 (b_{T\cdot C,X} - \beta_0) - \alpha_3 \beta_W W + \alpha_3(b_{T,C\cdot X} - \beta_C)C - \sum_{I=1}^{3} \alpha_3 \beta_{WCI} WCI + \alpha_3(b_{T,X\cdot C} - \beta_X).$$

With the coefficient signs assumed above, we see from (7) that an estimated version of (4) will probably systematically underestimate the labor supply of women who worked in the previous year (hence $W = 1$), especially for those who also have two or more children (hence $C > 2$ and either $WC2 = 1$ or $WC3 = 1$).

D. Prediction Accuracy, Proxy Effects, and Statistical Discrimination

Statistical discrimination occurs when an individual is passed over for employment or training or a promotion because of characteristics of some group the individual belongs to that are not characteristics of the individual.[7] The following passage illustrates the ''facts'' and logic that lead to statistical discrimination against women interested in working:

> Because of their dual role in the household and in the labour market, women traditionally have a shorter expected length of stay in the labour market. . . . In addition, their time in the labour market tends to be intermittent and subject to a considerable degree of uncertainty, thus creating rapid depreciation of their human capital. . . . For this reason it may be economically rational for females (or firms) to be reluctant to invest in female human capital formation that is labour market oriented (Gunderson and Riddell 1988, p. 452).

Blau and Ferber report that in interviews male executives repeatedly express the belief that women are less committed to their careers than men and are likely to quit their jobs when they have children. They go on to note:

> If such employer beliefs are simply incorrect or exaggerated . . . , actions based on them are clearly unfair and constitute labor market discrimination. . . . The situation is different . . . if the employer views are indeed correct on the average. . . . Yet, the consequences for individual women are far from satisfactory. A particular woman who would be as productive and as stable an employee as her male counterpart is denied employment or paid a lower wage. . . . Indeed, the practice of judging an individual on the basis of group characteristics rather than upon his or her own merits seems the

7. See Phelps (1972).

very essence of stereotyping and discrimination (Blau and Ferber 1986, p. 253).

An estimated version of Equation (4) will provide poor labor supply predictions for the women whose values of T are poorly approximated by the portion of (5) in braces. These women will have values for the OLS residual, $e_{T \cdot C,X}$, that are large in magnitude.

The labor supply of women with persistently large positive values of $e_{T \cdot C,X}$ will tend to be underestimated, period after period, if an estimated version of (4) is used for prediction. This could contribute to, or serve to justify, ongoing statistical discrimination against these women by employers. On the other hand, labor supply will tend to be overestimated for women with persistently large negative values of $e_{T \cdot C,X}$. Employers (and economic consultants and analysts) may react to the lower than expected realized labor supply of these women, and the higher than expected loss of training investments in them, by concluding that training any women is a mistake. The prevalence and seriousness of these prediction and statistical discrimination problems should be less the higher the value is for $R^2_{T \cdot C,X}$, the R^2 for the (unobservable) population regression in (5). This is also the situation, however, in which the behavioral response coefficient bias problem for Equation (4), discussed in subsection VB, is likely to be most serious.

E. An IV Solution to the Child Status Coefficient Bias Problem

Let Z denote a vector of observable variables that are correlated with C but not with $(\alpha_3 T + \varepsilon)$ in (1). Predicted values for C, denoted by $\hat{C}$, could be obtained from the auxiliary regression of C on Z; and C could be represented as

$$(8) \quad C = \hat{C} + e_{IV},$$

where e_{IV} is the residual term for the auxiliary regression.

If Z includes X, then e_{IV} will be orthogonal to X as well as $\hat{C}$ *by construction*. In this case, regression of h on $\hat{C}$ and X will yield consistent estimates of the parameters α_0, α_1, and α_2 of (1) and of

$$(9) \quad h = \alpha_0 + \alpha_1 \hat{C} + \alpha_2 X + [\alpha_1 e_{IV} + \alpha_3 T + \varepsilon],$$

where (9) is obtained by substituting (8) into (1). So long as appropriate instruments can be found for our simplified true model, the *IV* approach solves the coefficient bias problem associated with the direct inclusion of the child status variable in a model for h with T unobserved.

However, an estimated version of (9) will not provide unbiased predictions. Subtracting the systematic portion of the true reduced form Equation (3) from the expectation of the systematic portion of (9) yields the following expression for the prediction bias:

$$(10) \quad -\alpha_3 \beta_0 - \alpha_3 \beta_W W - \alpha_3 \beta_c C - \sum_{I=1}^{3} \alpha_3 \beta_{WCI} WCI - \alpha_3 \beta_X X.$$

We see from (10) that, given our assumptions about coefficient signs, an estimated version of Model (9) with the *IV* child status variable will tend to systematically underestimate the labor supply of women who worked in the previous year (with $W = 1$), especially for those who also have children (hence $C > 0$, and $WC1$ or $WC2$ or $WC3$ equals 1). In fact, comparing the bias expressions in (10) and (7), it is clear that the prediction bias problem could be worse for an estimated version of (9) with the IV child status variable than for (4) with the child status variable directly included. The terms $b_{T \cdot C,X}$, $b_{T,C \cdot X}$, and $b_{T,X \cdot C}$ in (4) [but not (9)] reflect the fact that some of the effects of T are picked up in (4) via C, with the consequence that these terms serve to reduce the prediction bias for (4), as evident from (7).

F. Reducing Prediction Bias Problems by Using Child Status and Lagged Labor Supply Variables as Proxies for Unobservable Tastes and Other Preconditions

For many purposes such as predicting the size and composition of the work force, predicting how women would be affected by changes in rules for entitlements such as pensions, and examining how the work behavior of women will affect family income inequality, what is needed are accurate predictions of the year to year work behavior of various demographic groups of women. In these situations, it may be sensible to deliberately use information about child status and previous work behavior as a way of capturing some of the effects of unobservable factors correlated with these variables.

Suppose that h is regressed on C, X, W, and WC. The estimated coefficients for this regression will be unbiased estimates of the coefficients of the population reduced form model reproduced here for convenience:

$$(3)\quad h = (\alpha_0 + \alpha_3\beta_0) + \alpha_3\beta_W W + (\alpha_1 + \alpha_3\beta_C)C + \sum_{I=1}^{3} \alpha_3\beta_{WCI}\, WCI + (\alpha_2 + \alpha_3\beta_X)X + [\alpha_3 v + \varepsilon].$$

If the objective is to obtain estimates of α_1 and α_2, then the terms

$$(11)\quad \alpha_3\beta_C \text{ and } \alpha_3\beta_X$$

in the coefficients for C and X in (3) will be biases. In this sense, (3) is an inferior empirical model to (9), the model with the instrumental child status variable. But there will be no prediction bias problem for an estimated version of (3) if (1) is true and v is random as has been assumed.[8]

VI. Conclusions

According to the usual explanations of economists as to why child status variables are included in models of female labor supply, the coefficients of these variables should be negative. By the selection and grouping of observations,

8. Notice that if the behavioral response coefficient bias problem were solved in (3) by substituting ($\hat{C} + e_{IV}$) from (8) for C throughout (3), we would be back in the situation of having an equation which fails to account for any of the differences among women in their unobservable values of T.

we control for the age, education, and marital status variables usually included in empirical models of female labor supply, and then examine the labor supply of women without and with children. Two measures of labor supply are used: weeks of work in a year and the propensity to be employed in a week. In line with the usual hypothesized effects of children, we find that married, childless women supply less labor than unmarried, childless women, and that labor supply is even lower for the married women with children. Also, examining cross-sectional evidence, we find for both the United States and Canada that it is married women, and especially married women with children, whose labor supply behavior has changed over the 1970/71–1980/81 period.

When the observations are also categorized by weeks of work in the previous year, we still find that labor supply (measured by the propensity to work in a week) declines moving from the unmarried, childless women to the married, childless women to the married women with children. We also find that labor supply rises steeply moving from the category of women with 0 weeks of work in the previous year, to those with 1–26 weeks, to those with 27–47, and finally to those who worked 48+ weeks in the previous year.

We go on to further group the observations for married women by the number of children ever born. Labor supply is found to fall moving from zero parity women to those with one child. But, to our surprise, for the groups of married women with children who worked in the previous year, we find that labor supply *rises* with parity. This is true even when we consider only married women with at least one child younger than six. One way this could happen is if the married women who had two or more children in the current year (and hence one or more children in the previous year), and who nevertheless worked in the previous year, have particularly strong tastes for work or needs for the income earned from working.

Finally, we explore the coefficient bias and the prediction bias implications of alternative estimation strategies in the context of a simplified true model that is consistent with the observed patterns in our cross-tabulations. We conclude that in situations where prediction accuracy is paramount, child status variables along with lagged labor supply variables should be explicitly and directly used as proxies for unobserved tastes and preconditions affecting female labor supply, as well as to capture any direct effects of children.

Data Appendix

Two basic sorts of information on work behavior are used in this study: employment status in the reference week and weeks of work in the previous calendar year. The reference week is the calendar week prior to the date on which respondents completed their census questionnaires or were interviewed by enumerators. Since the week of enumeration was not the same for all persons for any given census, the reference week is not a single week.

In the text, a person who was employed in the reference week is sometimes referred to as having "worked" then. Likewise, in the text a person is referred to as having "worked" in the previous year if the person had positive weeks of work then.

Details follow for the 1970 and 1980 U.S. censuses and the 1971 and 1981 Canadian censuses.

A. *1970 U.S. Census Data*

The employment status information for the reference week that is reported in the 1970 U.S. public use sample data is ascertained for persons 14 years of age and over from replies to several questions. These were: "Did this person work at any time last week (include part-time work such as Saturday job or helping without pay in family business or farm and active duty in the Armed Forces; exclude housework, school work, or volunteer work)? How many hours did he work last week (at all jobs)? Does this person have a job or business from which he was temporarily absent either because of illness, vacation, labor dispute, etc., or because he was on layoff last week? Has he been looking for work during the past four weeks, and if so, was there any reason why he could not take a job last week?" Civilians 14 years and over were counted as *employed* in the reference week if they were either "at work" (in other words, they did any work for pay or profit or worked without pay for 15 hours or more on a family farm or business) or reported being "with a job but not at work" (in other words, they were temporarily absent because of reasons such as illness, vacation, etc.).

The corresponding information on weeks worked was ascertained for persons 14 years of age and over who worked at all during the calendar year preceding the census from replies to two questions. These were: "Last year (1969) did this person work at all, even for a few days?" If yes, then "How many weeks did he work in 1969, either full-time or part-time?" Paid vacations, paid sick leave, unpaid work on a family farm or business, and military service are counted as weeks worked. The following time categories were presented: 13 weeks or less, 14 to 26 weeks, 27 to 39 weeks, 40 to 47 weeks, 48 to 49 weeks, and 50 to 52 weeks. It should be noted that the determination of weeks worked during the previous year was essentially independent of the determination of employment status in the reference week. See U.S. Department of Commerce (1972, pp. 151–52).

B. *1980 U.S. Census Data*

In the 1980 U.S. census, information about work behavior was collected for persons 16 years of age and older, rather than 14 years of age and older. This change has no impact on the analyses in this paper, however, which are for women 20–45 at the time of enumeration. For many persons the reference week for the 1980 U.S. census is the last week of March, 1980.

A second change from 1970 is that, on the public use sample tapes, actual weeks worked in the previous year (1 to 52) are given rather than categorical information. We used this information to determine the category for weeks of work in the previous year (0, 1 to 26, 27 to 47, and 48+) for each observation.

The definitions for the categories were chosen to be compatible with the weeks of work information provided in the 1970 U.S. census.

For further details see U.S. Department of Commerce (1983, pp. K–25 to K–26 and K-54).

C. 1971 Canadian Census Data: Family File

The employment status information for the reference week that is reported in the 1971 Canadian public use sample data is ascertained for persons 15 years and over at the time of enumeration from replies to five questions. Respondents were asked for, or about: "Hours worked last week for pay or profit; Hours helped last week without pay in family farm or business; Looked for work last week; With job but on temporary lay off last week; With job last week but absent from work." Respondents were counted as employed in the reference week if they reported that they worked for pay or profit, or in unpaid family work, or had a job from which they were temporarily absent because of illness, vacation, strike, or other such reasons, with a few exceptions. Persons who had jobs but did not work during the reference week were counted as unemployed (and hence not employed) if they reported being on temporary lay off or that they looked for work. Female farm workers working less than 20 hours in the week on a family farm or in a family business without pay were also excluded, as were employed inmates of institutions.

For most persons, the reference week was the week of May 24 to 31, 1971.

For the previous calendar year, those who worked for even a few hours are counted as having "worked" in the previous year. The number of weeks of work in the previous year includes weeks of paid vacation or sick leave or paid absence on training courses. Self-employment weeks, weeks of unpaid work on a family farm or for a family business, and work for payment "in kind" in a nonfamily enterprise are counted along with weeks of work for "pay or profit." Information is available for the following categories: 1 to 13, 14 to 26, 27 to 39, 40 to 48, and 49 to 52 weeks. Thus, the figures shown in Tables 1 and 3–5 for Canada 1971 under the labels of 27 to 47 and 48+ weeks worked in the previous year are actually for 27 to 48 and 49+ weeks. Coleman and Pencavel (1992) show separate frequency distributions, by single weeks (1–52), for white women and for black women in the United States, as reported in the 1940 and in the 1980 U.S. censuses. These show large spikes at 52 weeks and lower spikes at 50 and 40 weeks (and at 26 weeks for 1940 but not 1980), but relatively few observations at 48 weeks. Thus the consequences of the problems for our 27 to 47 and 48+ weeks in the previous year categories for Canada 1971 are not believed to be serious.

For further details, see Statistics Canada (1975, pp. 20–22 and p. 33).

D. 1981 Canadian Census Data: Household/Family File

All of the concepts are the same as for the 1971 Canadian census. However, the reference week for most persons is the week before June 3, 1981. Also, weeks in the previous year are given in single weeks (1 to 52). Hence we were able to sort observations properly into our 0, 1 to 26, 27 to 47, and 48+ categories for weeks of work in the previous year.

For further details, see Statistics Canada (1985).

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