Popul Res Policy Rev (2007) 26:235–257
DOI 10.1007/s11113-007-9029-6

Effects of child support and welfare policies on nonmarital teenage childbearing and motherhood

Lingxin Hao · Nan M. Astone · Andrew J. Cherlin

Received: 27 July 2005 / Accepted: 14 April 2006 / Published online: 14 March 2007

Abstract This paper is an assessment of the impact of child support enforcement and welfare policies on nonmarital teenage childbearing and motherhood. We derive four hypotheses about the effects of policies on nonmarital teenage childbearing and motherhood. We propose that teenage motherhood and school enrollment are joint decisions for teenage girls. Based on individual trajectories during ages 12–19, our analysis uses an event history model for nonmarital teenage childbearing and a dynamic model of motherhood that is jointly determined with school enrollment. We find some evidence that child support policies indirectly reduce teen motherhood by increasing the probability of school enrollment, which, in turn, reduces the probability of teen motherhood. This finding suggests that welfare offices may wish to place greater weight on outreach programs that inform more teenagers of the existence of strong child support enforcement measures. Such programs might reduce nonmarital teen motherhood further and thus reduce the need for welfare support and child support enforcement in the long run.

Keywords Child support policy · Welfare policy · Teenage childbearing · Teenage motherhood

Introduction

A prominent theme in the discourse about the rise in the nonmarital birth ratio in the United States is the role of lax child support enforcement and generous welfare policy in creating unintended incentives for couples to have

Lx. Hao (✉) · N. M. Astone · A. J. Cherlin
Department of Sociology, Johns Hopkins University, Baltimore, MD 21218, USA
e-mail: hao@jhu.edu

children outside marriage. A series of policy initiatives to strengthen the enforcement of child support by noncustodial parents was one response to this discourse (Garfinkel et al. 1998). Another response was welfare reform, the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA). The latter had many public policy goals, but perhaps none was so greatly noted as the reduction of early nonmarital childbearing, which was identified in the preamble to the legislation itself as a primary objective.

The underlying ideas behind these policy changes are as follows: If (a) young people expect that, as noncustodial parents, their obligation to economically support their children will be strictly enforced and (b) young people expect that, as custodial parents, their reliance on public transfers to support their children will be tightly constrained, then, young people will, on the one hand, adjust their sexual, contraceptive, abortion, and marriage behaviors in ways that will reduce the likelihood of nonmarital births, and on the other hand, obtain sufficient schooling to increase their future capacity to support a family.

According to this view, stringent child support policies provide a disincentive for men to father children outside of marriage. Similarly, the former aid program, Aid to Families with Dependent Children (AFDC), by guaranteeing custodial parents a source of support until their youngest child reached the age of 18, provided an incentive for couples to remain unmarried after discovering an unplanned pregnancy. Policy initiatives to strengthen the enforcement of child support mandate the establishment of paternity and to provide services to locate the father in order to collect child support payments. PRWORA, by replacing AFDC with a program called Temporary Assistance to Needy Families (TANF), mandates that states implement time limits on the receipt of TANF benefits, work requirements for most TANF recipients, and sanctions on TANF recipients for noncompliance with the rules. It also requires teenage mothers under the age of 18 to live with their parents and remain enrolled in high school if they have not yet graduated in order to obtain welfare benefits. The increased stringency of child support policies and TANF was intended to ensure that children are eventually economically supported by their parents, not by the government through public transfers. Its proponents hope that, knowing their greater responsibility, young people will delay parenthood until they have acquired sufficient human capital to support a family adequately and are prepared to marry.

It is crucial to assess the effectiveness of these policies on the behaviors they were intended to influence. If we find that they have indeed reduced nonmarital teenage childbearing, then the policies can be judged a success and their underlying rationale will receive greater credence. If, however, nonmarital teenage childbearing has not declined under these policies, then their continued application must be called into question.

In this paper, we examine whether or not this intended effect of child support policies and PRWORA has occurred in the short term. To do this, we take advantage of two notable aspects of child support and welfare policies in the 1990s. First, states are given wide latitude in designing and implementing their specific child support and welfare policies, so there is extensive variation across states in how stringent state child support policies and public assistance programs are. Second, during the years immediately preceding PRWORA, 29 states granted waivers by the federal government implemented many of the policies PRWORA later adopted. In addition, during the early 1990s there were large changes in state child support policies. Thus, for most of the 1990s, residents of different states have confronted a set of incentives and disincentives to their behavior emanating from child support and welfare policies that are differentially *stringent* (e.g., requiring teenage recipients to live with their parents and enroll in school). This situation facilitates research on the effects of state policies since researchers may compare the behavior of similar people who live in different policy environments.

We use longitudinal data from the National Longitudinal Survey of Youth 1997 cohort (NLSY97) combined with state-level data sources. Our analysis has several features: (a) we examine nonmarital fertility behavior of adolescents rather than all reproductive women to reveal the policy effects for this less-investigated age group; (b) we examine the policy effects on teen motherhood and school enrollment jointly to reflect the fact that these decisions are made jointly, (c) we control for other state variables that may affect early nonmarital childbearing and schooling to separate out the effects of welfare and child support policies; and (d) we control for unobserved heterogeneity across states that may confound the effects of welfare and child support policies.

Background

Child support policies and their effects for parents and children

Since most noncustodial parents are fathers, the literature examining the effects of child support enforcement focuses on the behavior of fathers. The most obvious question regarding the effectiveness of stringent child support policies is whether or not they result in increased child support payments. In general they appear to have such an effect (Freeman and Waldfogel 1998). But Bloom et al. (1998) point out that this is a more complicated question than it first appears. In the aggregate, children benefit from both the payment of child support from their biological father and their mother's remarriage to another man. Their data indicates that while enforcement policies do appear to increase child support, they decrease remarriage rates, and it is not clear what the net overall gain to children is from strict enforcement. With respect to the fathers' economic status, Meyer (1998), using simulations, estimates that full payment of child support would have no effect on the poverty status of lower-income fathers, although its impacts on higher-income fathers are greater (because they typically have higher obligations). Freeman and Waldfogel (1998) find very modest positive effects of child support enforcement on male labor supply, debunking the hypothesis that men will respond to stricter enforcement by working less.

A number of recent studies have evaluated the effects of child support enforcement and welfare on nonmarital childbearing of reproductive-age women. Using aggregate state-level data, Garfinkel et al. (2003) find a robust effect of child support enforcement, which reduces nonmarital birth rates among women aged 15–44 during the 1980–1996 pre-welfare reform period. They find a weaker effect of welfare benefits on nonmarital birth rates. Using individual data, Huang (2001) examines the role of strict child support legislation and high child support expenditure and finds that these policy measures increase the likelihood of marital births and decrease the likelihood of nonmarital births among women aged 14–40 during the 1979–1998 period. Plotnick et al. (2005) study the childbearing behavior of women aged 15–44 and find that women living in states with more effective child support enforcement are less likely to bear children when unmarried. These studies provide evidence for the importance of child support policies for reproductive-age women. Our study differs from these recent studies because our study focuses on adolescents' childbearing decisions, isolates the effects of child support policy and welfare reform policy from unobserved state characteristics, examines the years 1992–2001, and uses a dynamic (transition) model.

Effects of welfare and welfare reform on nonmarital births

Research on the incentive effects of welfare benefits on nonmarital childbearing suggests that the level of welfare benefits before waivers has a real but modest effect on nonmarital fertility (Moffitt 1998). Foster and Hoffman (2001) find that welfare policy appears to affect nonmarital fertility more among older women than among teenagers and that any positive effect of AFDC policies on nonmarital childbearing is eliminated when unobserved heterogeneity across states was controlled. While Blank (2002) suggests that the sweeping changes under PRWORA could produce a larger effect of welfare reform on fertility, more recent research using experimental and nonexperimental designs has not yet provided a consensus. Fein (1999) finds little effect on fertility of a strong mandatory work activities program. Quint et al. (1997) find an increase in the teenage pregnancy rate under a program that provided educational and job assistance to teenage welfare mothers. Hao and Cherlin (2004) examine the fertility behavior of adolescents aged 14–16 by comparing a cohort before and a cohort after the implementation of state welfare reform using a difference-in-differences method. They find that welfare reform has not reduced teenage pregnancies and births. Kaestner et al. (2003) and Offner (2003) find that welfare reform had a significant, small negative effect on the fertility of young women in low-income families.

The sociological literature on teenage childbearing has identified individual/family factors that increase the risk of teenage pregnancy, birth, and school dropout. These include family background factors such as race/ethnicity, family income, parental education, family structure, and number of siblings (Jones et al. 1999; Hardy et al. 1998; Wu 1996; Wu and Martinson 1993). Parental welfare participation is another micro-factor worthy of attention. Findings from studies examining how maternal welfare receipt affects daughters' fertility are contradictory. Gottschalk (1990, 1992) finds that maternal welfare receipt is a risk factor for daughters' premarital births, whereas Haveman and Wolfe (1994) find no association between maternal welfare and daughters' early fertility.

Conceptual framework

As reviewed above, the nonmarital childbearing literature considers factors at the individual/family level and the role of policy in determining nonmarital childbearing. Although we follow the usual approach to examining individual/family factors, our modified approach is to focus on the role of policies by assessing the impact of the policy environment where a youth is situated. A life course perspective allows us to build the conceptual model for this policy focus.

One assumption of the life course perspective, the linkage between individual life course events and larger structural social changes, provides a basic rationale for our conceptual model. If a state makes a significant shift in child support and welfare policies that directly alters the context within which a young couple will rear a child, young people residing in the state should change their preferences regarding nonmarital fertility behavior. Specifically, if child support is strictly enforced, a greater proportion of young men should avoid fathering children outside of marriage in order to avoid the substantial long-term economic obligations to support the child. Consequently, girls should be at lower risks of conceiving and giving birth. Therefore, we would expect a depressing effect of strict child support policies on girls' likelihood of having children outside of marriage. Likewise, if welfare policies are more stringent and public assistance is tightly constrained, more girls should depress nonmarital childbearing to avoid becoming the sole supporter of a child in situations where the father has low support capacity.

Another assumption of the life course perspective—namely, that life events are interdependent—provides an additional rationale for our conceptual framework to study policy effects. The policy environment could affect girls' school enrollment in two ways. First, welfare policies mandate that unmarried mothers under age 18 who receive welfare attend school and live with their parents, who help enforce school enrollment. School enrollment would go up with the presence of these mandates. Second, the costs of nonmarital teen motherhood (i.e., the opportunity costs of not being enrolled in school) go up with welfare reform because welfare is no longer timeless. Thus, welfare reform should increase school enrollment. Third, the costs of children increase with stricter child support for young men so that young girls are less distracted from schooling, thereby increasing the likelihood of young girls' school enrollment. Consequent to these two policy effects is an increase in the probability of school enrollment, which, in turn, may further deter early nonmarital motherhood because it is documented that the birth rates of high school students are lower than those of girls of similar age who are not enrolled (Astone and Upchurch 1994; Jones et al. 1999). Therefore schooling and motherhood among teen girls are interrelated and the decisions regarding them are made jointly. To capture the indirect policy effects of welfare policies on nonmarital childbearing through school enrollment, we model the interdependence of the two behaviors and their joint decision over the life course of adolescence.

In sum, our central task is to test the following four hypotheses. First, child support policies depress teenage girls' nonmarital childbearing and childrearing because the risk of pregnancy/childbirth is lowered by young men's adjusted sexual behavior. Second, welfare policies depress teenage girls' nonmarital childbearing and childrearing because they require enrollment in school and living with parents and the anticipated public assistance is limited. Third, child support policies and welfare policies can promote school enrollment because of the heightened opportunity costs of not attending school and the heightened costs of children. Fourth, child support and welfare policies suppress teenage childbearing and childrearing indirectly though their promoting effects on teenage girls' school enrollment.

In our consideration of the policy effects on nonmarital childbearing/motherhood and school enrollment, we pay special attention to why policies can affect adolescents while the target population of the policies is adults. First, child support policies affect young men whose behavior is an integral part of teenage parenthood; this is often ignored. In contrast, welfare policies address custodial parents and thus they should affect girls' fertility behavior. Second, some adolescents live in households that depend on welfare benefits for income or have characteristics that make welfare receipt a likely option, such as being headed by a single parent or being poor. The decisions of adolescents in this latter group may be more responsive than others to changes in the incentives and disincentives of child support and welfare policies, since they are more likely than others to regard these policies as something that could affect their lives. Therefore we examine the differential effects of child support and welfare policies for poor and non-poor adolescents and expect stringent policies to have a significantly stronger impact on poor than non-poor youths.

Data and methods

The strictness of child support enforcement policies varies by year from 1992 to 1998 and by state. Since 1998, most child support policies are enforced within the welfare reform legislation and there is little variation over the years. Because the topic of study is early nonmarital fertility, we searched for features of child support policy most relevant to young men's and women's fertility behavior, for example, the extent to which paternity establishment is enforced and the father is located to enforce child support payments. We created two indexes from four indicators: (a) the paternity establishment index is made up of the proportion of all nonmarital births that have established paternity and the per birth expenditure in laboratory paternity establishment, and (b) the locator service index is made up of the number of locator services to trace fathers with or without social security numbers, normalized by the number of nonmarital births. The data are drawn from Department of Health and Human Services (DHHS).[1]

We contend that the minor mandates of welfare policy (attending school and living with parents) are most relevant to teenagers' fertility decisions. Based on these two mandates, we created an index that varies by state from 1992 to 1999. Since 1999 there is no state or time variation in these mandates. This index is highly correlated (.85) with the stringency index used in Hao et al. (2004), which was based on state and time variations in work requirements, time limits, sanctions, work exemption based on child age, and the two minor mandates. Therefore, in this paper we include only the minor mandate index to capture stringency of welfare policy prior to and post the welfare reform law. The data are drawn from the Office of Assistant Secretary for Planning and Evaluation, Department of Health and Human Services DHHS.[2]

We also consider other state policies that may confound child support and welfare policies in influencing adolescents' fertility behavior. For example, Joyce and his colleagues have demonstrated over the years that abortion policy affects fertility behavior (Joyce and Kaestner 1996; Joyce et al. 1998). We collected abortion policy information from the National Abortion and Reproductive Rights Action League. We created an index for state abortion policies based on the minimum waiting hours for an abortion to be performed, the requirement of parental consent, and the conditions under which public funds could be used to support abortion, such as life endangerment, rape or incest, health concerns, and others. Moreover, schooling and fertility decisions may be affected by local labor market conditions, which we measure using the state- and year-variant unemployment rates obtained from the Bureau of Labor Statistics. These factors affect adolescents' access to abortion and perception of their labor market options, which may affect decisions about other realms of life such as motherhood and schooling.

Other important factors are the political and social climates of the state, which are difficult to observe. These unobserved state characteristics may correlate with the measured state characteristics, resulting in biased estimates for the state policy variables (see Foster and Hoffman 2001). We control for

---

[1] The websites are http://www.acf.dhhs.gov/programs/cse/rpt/21t/st26.pdf, http://www.acf.dhhs.gov/programs/cse/rpt/22t/TABLE40.htm, and http://www.acf.dhhs.gov/programs/cse/rpt/annrpt23/tables/TABLE40.htm.

[2] The website is http://aspe.hhs.gov/hsp/Waiver-Policies2000. Data were available on the approval date and implementation date for various welfare programs in 50 states over the years 1992–2000. We used the implementation date to construct our variable.

unobserved, time-invariant state characteristics by including in the model a set of dummy variables for each of the states (minus 1).

Following the literature on early childbearing we include in our models individual characteristics (age) and family background variables (race/ethnicity, family structure, family income, parental education, parental AFDC status,[3] number of siblings, and urban residence), which come from the NLSY97, described in detail below. We estimate the effects of stringent child support and welfare policies on nonmarital teenage childbearing net of these individual/family factors and the other state characteristics we described above.

The NLSY97 is a nationally representative sample of 8,984 individuals age 12–16 as of December 31, 1996, with an oversample of black and Hispanic youth. After the first interview in 1997, 8,386 respondents received the second interview during 1998–1999, 8,209 received the third interview during 1999–2000, and 8,081 received the fourth interview during 2000–2001. This study uses the panel data of Wave 1 to Wave 4 on female respondents. The retention rate at Wave 4 is 89.9%. Our approach to the attrition problem is to retain all observations of female respondents age 12–19 over the course of the panel, resulting in 4,348 female respondents. Rather than using only those respondents with all four waves of data, our approach can reduce the potential selection bias by attrition.

In addition to the prospective information on fertility, pregnancy, and enrollment from the four interviews, retrospective information on these histories were collected from all respondents, so we can construct three microhistories for each respondent. The fertility history identifies those who had given birth outside marriage during their teens. The event history for each female respondent starts from exact age 12 (age 144 in months) until the month when the first birth occurs before exact age 20 or the end of observation (age 239 months or earlier), whichever occurs first. We left-censored four girls who had a birth before age 12. We right-censored marriage before a teen birth. That is, the history ends one month before marriage without a birth. We use the fertility history data to study the event of first nonmarital teenage birth. Table 1 shows the distribution of nonmarital teenage births by age. The table shows that during the age of 12, out of 4,348 girls, two had births and 18 were no longer observed beyond age 12. The birth rate is .05% at age 12 and increases over age to reach 5.3% during age 19. Altogether, we observed 535 nonmarital teenage births; 351 occurred during ages 12–17 and 181 during ages 18 and 19.

We built a *motherhood* history that combines pregnancy and birth data. Motherhood is defined as a state of bearing or rearing a child. A pregnancy can end in miscarriage, abortion, a still birth, or a live birth. After a child is born, the mother can raise the child on her own or make an adoption plan for her child. In this history, the state of being in or out of motherhood in each

[3] Parents' AFDC status is available only before 1997. We distinguish those who received AFDC before 1992 but not after 1992 and those who received AFDC 1992–1996, regardless of whether receiving AFDC before 1992. We did not include the youth's AFDC/TANF status because only very few youths ever received AFDC.

Table 1 Distribution of nonmarital teenage births by age

Age	At risk	Birth	Rate (%)	End of observation
12	4,348	2	.05	18
13	4,328	4	.09	13
14	4,311	33	.77	33
15	4,245	66	1.55	69
16	4,110	120	2.92	790
17	3,200	126	3.94	855
18	2,219	114	5.14	785
19	1,320	70	5.30	1,250
Total		535		

Data sources: NLSY97 waves 1–4 and retrospective fertility history

month from age 12 to age 19 is recorded. As for the fertility event history, we left-censored four girls and right-censored on pre-birth marriage for the motherhood history. Different from the fertility event history, the motherhood history does not end in a birth but continues, allowing movements into and out of motherhood.

A third history is that of school enrollment. The state of being enrolled in formal schools in each month from age 12 to age 19 is recorded. Movements into and out of the enrolled state are allowed. The enrollment history right-censors high school graduation and college attendance. Pooling motherhood history and enrollment history, we study the interdependence of motherhood and schooling. The combined enrollment and motherhood histories thus right-censor at pre-birth marriage, high school graduation, college attendance, or end of observation before exact age 20. Table 2 examines the percentage distribution of secondary school enrollment and nonmarital teenage motherhood by age. Among the 43,070 person-months we observed during age 12, 99.38% were enrolled in school, .09% were enrolled in school and in motherhood, .53% were not enrolled in school nor in motherhood, and no one was

Table 2 Percentage distribution of school enrollment and nonmarital teenage motherhood by age

Age	Enrolled, not in motherhood	Enrolled, in motherhood	Not enrolled, not in motherhood	Not enrolled, in motherhood	Person-months
12	99.38	.09	.53	.00	43,070
13	98.75	.45	.77	.03	51,401
14	97.21	1.43	1.26	.10	51,069
15	94.87	3.22	1.44	.46	50,232
16	91.08	5.84	2.12	.96	44,538
17	86.65	8.70	2.67	1.98	31,375
18	74.33	12.80	6.53	6.34	9,874
19	42.18	15.49	17.94	24.39	1,956

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text)

Note: Motherhood is defined as being pregnant or raising a child outside of marriage and before high school graduation not enrolled and in motherhood. The distribution changes noticeably over age. For example, during age 17, 8.7% were enrolled while being mothers and 1.98% were not enrolled while being mothers.

We use a discrete-time event history model to test our first two hypotheses about the effects of child support welfare policies on nonmarital teenage childbirths. The discrete time model incorporates time-variant variables, including our key explanatory variables of child support and welfare policies in the state of residence. Time-variant variables include the paternity establishment and locator service indexes of child support policies, the minor mandate index of welfare policies, the abortion policy index, and state unemployment rates at the state level and family structure, income, number of siblings, and urban residence at the individual level. Time-invariant variables include a set of dummy variables for each of the 36 states with a sufficient number of person-months and race/ethnicity, parental education, and parental AFDC receipt. Table 3 presents the descriptive statistics of time-variant and invariant variables. Panel A is for individual/family variables measured at age 14 for all girls, girls without and with teen births. Panel A shows that, compared with their no-birth counterparts, girls with births are more likely to be black or Hispanic, living in single-mother families or other types of family, living in poverty, have low parental education, have parents receiving AFDC, and have a large number of siblings. Panel B shows the mean of state-level variables for all person-months, those person-months by age 18 and those during ages 18 and 19. Note the different mean of these variables measured during the younger versus older teen ages of our respondents: for example, unemployment rate is lower, abortion policy is less stringent, paternity establishment is tougher, and the minor mandate index is stricter during the later teens than the earlier teens. In addition, the minor mandate index has no variation during old teen ages. Panel C shows the correlation coefficients among state-level variables, which are moderate, with the highest being –.55 between unemployment rates and the minor mandate index.

We use a model for the interdependence of two endogenous processes to test the four hypotheses about the effects of child support policies and welfare policies on nonmarital teenage motherhood and school enrollment. The analysis takes into account the potential reverse causal relationship between school enrollment and teenage motherhood by specifying their cross-lagged relationships. The enrollment history and motherhood history enable us to sort out the causal order between the two endogenous variables through constructing proper feedback effects for given individuals as they change status with respect to the two endogenous variables (Finkel 1995; Hsiao 2003). We adopt a model advanced by Yamaguchi (1990) that permits a reciprocal lagged causal analysis of the two endogenous processes and the simultaneous analysis of transitions between the two states of each process. Hao (1997) developed a method that uses a multinomial logit estimator and links multinomial logit parameters to the parameters that fit the two endogenous processes. The dependent variable of the multinomial logit estimator is the cross-classification of enrollment and motherhood, resulting in the four states

Table 3 Descriptive statistics of variables used in analysis

Variable	All girls		Girls w/o birth		Girls w/ birth	
	Mean	SD	Mean	SD	Mean	SD
(a) Individual/family variables at age 14						
Black	.27	.44	.24	.43	.44	.50
Hispanic	.21	.41	.20	.40	.26	.44
Other race	.04	.19	.04	.20	.03	.18
Stepfamily	.13	.34	.13	.34	.14	.34
Single-mother	.29	.46	.27	.44	.46	.50
Single-father	.03	.18	.04	.18	.03	.17
Other family	.06	.23	.05	.22	.12	.33
Income-to-needs	2.93	2.76	3.09	2.85	1.79	1.62
Poverty	.26	.44	.23	.42	.48	.50
Parental education	13.16	2.94	13.32	2.95	12.03	2.59
Parental AFDC before 1992	.12	.32	.12	.32	.15	.36
Parental AFDC 1992–97	.16	.36	.14	.34	.31	.46
# of siblings	1.64	1.35	1.62	1.31	1.81	1.58
Urban residence	.76	.43	.75	.43	.79	.41
# of individuals	4,348		3,813		535	

(b) State-level variables

Variable	All person-months		Person-months by 18		Person-months post 18	
	Mean	SD	Mean	SD	Mean	SD
1. Unemployment rate	5.04	1.33	5.15	1.34	4.15	.88
2. Abortion policy index	–.14	.71	–.15	.71	–.10	.68
3. Paternity establishment index	–.08	.71	–.09	.71	.02	.68
4. Locator service index	3.32	3.30	3.33	3.33	3.23	3.06
5. Minor mandate index	.31	.89	.23	.91	1.01	.00
# of person-months	315,911		282,971		32,940	

(c) State-level variable correlation

	1.	2.	3.	4.
2.	–.2844			
3.	–.4204	.1244		
4.	.0829	–.2415	.1406	
5.	–.5502	.0872	.1647	.0298

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text)

described in Table 2. The effects of child support and welfare policies are specified in the model to test hypotheses 1–4 regarding teenage motherhood and school enrollment. Lagged endogenous variables are included in the model to examine the cross-lagged effects between the two processes proposed in Hypothesis 4. This method allows us to specify the interdependence of two endogenous processes using existing statistical packages (Stata is used in this study). The converted estimates are simple linear combinations of the multinomial logit estimates; standard errors of the converted estimates can also be obtained through linear combinations of elements in the variance-covariance matrix of the multinomial logit estimates (see the Technical Appendix for a brief description).

To address the concern about repeated observations of an individual in our analysis, we use the Huber correction to adjust the standard errors of estimates (Huber 1967). This correction produces robust estimates of standard errors of estimates correcting for the dependence of repeated observations of the same individual over time. The statistical tests are then based on the number of individuals rather than the number of person-months.

Results

Nonmarital teenage childbearing

Tables 4 and 5 together present the results from the event history model of nonmarital teenage childbearing. From the same model estimation, the results

Table 4 Effects of individual/family characteristics on nonmarital teenage childbearing

Variable	Young teen w/o state FE	Young teen w/ state FE	All teen w/o state FE	All teen w/ state FE
Age (in months)	.06**	.06**	.04**	.04**
	(.00)	(.00)	(.00)	(.00)
Black	.46**	.47**	.48**	.46**
	(.15)	(.17)	(.12)	(.13)
Hispanic	.45*	.41*	.49**	.41*
	(.18)	(.20)	(.15)	(.16)
Other race	.18	.23	.30	.30
	(.34)	(.35)	(.26)	(.27)
Stepfamily	.74**	.73**	.56**	.57**
	(.22)	(.23)	(.18)	(.18)
Single-mother	.92**	.88**	.74**	.71**
	(.17)	(.18)	(.14)	(.14)
Single-father	.68	.66	.50	.53
	(.36)	(.37)	(.30)	(.31)
Other family	1.77**	1.73**	1.46**	1.44**
	(.19)	(.19)	(.14)	(.15)
Income-to-needs	–.27**	–.28**	–.18**	–.19**
	(.08)	(.08)	(.05)	(.05)
Parental education	–.07**	–.07**	–.06**	–.05**
	(.02)	(.02)	(.02)	(.02)
Parental AFDC before 1992	.16	.17	.23	.22
	(.17)	(.18)	(.14)	(.14)
Parental AFDC 1992–97	.26	.26	.39**	.40**
	(.15)	(.15)	(.13)	(.13)
# siblings	.08*	.09*	.06	.05
	(.04)	(.04)	(.03)	(.03)
Urban residence	.10	.01	.09	.01
	(.14)	(.16)	(.12)	(.13)

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text)

Note: Presented are estimates from a logistic regression model (standard errors in parentheses). The model includes state-level variables which are presented in Table 4

** $p < .01$, * $p < .05$, ^ $p < .10$ (two-tail test)

Table 5 Effects of state-level variables on nonmarital teenage childbearing

Variable	Young teen w/o state FE	Young teen w/ state FE	All teen w/o state FE	All teen w/ state FE
Unemployment rate	–.03	.00	–.01	.01
	(.07)	(.11)	(.05)	(.09)
Abortion policy index	.21*	–.01	.18**	.28
	(.08)	(.28)	(.07)	(.25)
Paternity establishment*non-poor	.03	.02	.13	.04
	(.12)	(.23)	(.09)	(.19)
Paternity establishment*poor	.22	.19	.18	.07
	(.12)	(.21)	(.09)	(.17)
Locator service*non-poor	–.04	.05	–.04*	.03
	(.03)	(.05)	(.02)	(.04)
Locator service*poor	–.07*	.01	–.04	.02
	(.03)	(.04)	(.02)	(.04)
Minor mandate index*non-poor	–.09	–.06	.14	.16
	(.14)	(.15)	(.12)	(.13)
Minor mandate index*poor	–.06	–.02	.15	.19
	(.14)	(.15)	(.13)	(.14)

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text)

Note: Presented are estimates from a logistic regression model (standard errors in parentheses). The model includes individual/family variables which are presented in Table 3

** $p < .01$, * $p < .05$, ^ $p < .10$ (two-tail test)

about individual/family variables are presented in Table 4 and the results about state variables are presented in Table 5. Previous research suggests that social contexts beyond the family play an increasing role in adolescents' life course. To see whether young girls are more heavily influenced by family background factors when making decision about teen births, we estimate the event history up to the month before exact age 18 for young teen girls and then estimate the full event history up to the month before exact age 20 for all teen girls. For each sample, we estimate two models: one specifies individual/family factors and state variables without a set of dummy variables for the 36 states, which are called a model without state fixed-effects (FE), and the other with state FE.

Table 4 presents the effects of individual/family factors on teen fertility. Comparing the results between the model without and with state FE, we observe that the coefficients remain almost identical for both young teens and all teens. This indicates that individual/family factors are uncorrelated with unobserved state characteristics. By comparing the results for young teens with those for all teens, we can infer the difference in coefficients for young teens versus old teens. The effects of individual/family factors on teen births are in the same direction as found in previous research, but the strengths and significance levels differ between young teens and all teens. Teen births increase with age, more strongly for young teens than for all teens. Black and Hispanic girls are at a higher risk of having a teen birth outside marriage than white girls and the effects are similar for young teens and all teens. Living with stepparents, single mothers, or nonparents increases the risk of having a teen birth, more so for young teens than for all teens. Higher income-to-needs ratios and parental education reduce the likelihood of having a teen birth, which is stronger for young teens than for all teens. While parental AFDC has no significant effect for young teens, recent parental AFDC received during 1992–1997 increases the probability of having a teen birth among all teens. These findings uniformly speak to the greater importance of family background factors influencing young teen girls' fertility behavior than older teen girls'. This is reasonable because as girls grow older, influences beyond the family such as from schools, teachers, peers, and neighborhood loom more important in girls' decisions on nonmarital childbearing. Parental welfare is an exception. It is possible that older teenagers understand more than young teenagers about the meaning of the state's financial support for their own family and they are able to generalize the state's support for a nonmarital birth, and thus take the incentives of welfare benefits into their teen birth decisions.

Table 5 presents the estimates for state variables from the same event history model for nonmarital teenage childbirths. We estimate the marginal policy effects for poor youth, (i.e., the effect and its standard error are for the poor subpopulation) by interacting a policy variable with the indicator for being poor. In parallel, we estimate the marginal policy effects for non-poor youths, (i.e., the effect and its standard error are for the non-poor subpopulation) by interacting a policy variable with the indicator for being non-poor. Comparing the coefficients between the two models without and with state FE, we see some important changes. The following interpretations focus on these changes.

State unemployment does not enter either young or old teen girls' nonmarital birth decisions. Stricter state abortion policies appear to increase nonmarital births among young and old teenagers in the model without state fixed-effects. Controlling for state fixed-effects alters the results, however. The effect reduces to almost zero for young teens, implying a bias in the estimate due to the correlation between unobserved state characteristics and state abortion policies. For all teens, however, the effect increases from .18 to .28 while the standard error also increases, making the estimate imprecise.

Among our key policy variables, the locator service index reduces the risk of birth among poor young teens if state FE is not controlled. When the state FE is controlled, however, this negative effect reduces to near zero. The estimates for all teens show that the locator service reduces the risk of birth among non-poor girls under the model without state FE, and this effect reduces moderately and becomes insignificant after state FE is controlled. We do not find any effect of the minor mandate index for young or all teens, without or with state FE being controlled. These results suggest that either child support policies or welfare policies play a significant effect on teenage girls' decision on births outside of marriage.[4] We should see these results in light of the sweeping change in welfare policies after the welfare reform law

[4] We performed a parallel analysis for young men aged 14–21 using the NLSY97 data based on the consideration that men are about 2 years older than women in sexual relationships. The results mirror what we found for teenage girls.

was fully implemented in 1998–1999 when most of our respondents were in the latter stage of their adolescent life course. The smaller variations in policy variables during the later stage of the life course may not allow us to identify significant policy effects on nonmarital teenage births.

School enrollment and nonmarital teenage motherhood

One limitation of the teen birth analysis is its assumption that the decision regarding teen births is independent of school enrollment decision. As we developed in the conceptual model, decisions on teen motherhood are directly related to decisions on school enrollment among teen girls. The teen birth model then ignores a possible route of policy influence. In Hypotheses 3 and 4 we expect that child support and welfare policies increase the likelihood of school enrollment, which in turn depresses nonmarital teenage motherhood.

To test Hypotheses 3 and 4 and retest Hypotheses 1 and 2, our next analysis is to use a model of the interdependence between nonmarital teenage motherhood and school enrollment. Recall that motherhood is defined as bearing or raising a child and enrollment is in schools. A girl can enter motherhood by becoming pregnant and exit motherhood by ending the pregnancy without a child or adopting out the child. She can also enter and exit the state of being enrolled in school. Our model simultaneously specifies three things: (a) the effects of child support policies and welfare policies on teenage motherhood and school enrollment controlling for individual/family factors; (b) the cross-lagged effects of the two endogenous variables, i.e., lagged motherhood affects current enrollment and likewise lagged enrollment affects current motherhood; and (c) state dependence, i.e., the current state of motherhood (or enrollment) depends on the lagged state of motherhood (or enrollment) controlling for covariates and cross-lagged effects. The time unit is a month. With these specifications, our model is a monthly transition model of two interdependent processes.

Tables 6 and 7 present the coefficients for the process of school enrollment and the process of motherhood, converted from estimates of 4-category multinomial logit models (see Technical Appendix). The converted coefficients represent the change in the logarithm of the probability of a state in each of the two interdependent processes caused by one unit change in the explanatory variable. For example, an increase of one year of parental schooling leads to an increase of .03 in the log enrollment probability and a decrease of .03 in the log motherhood probability.

Table 6 presents the effects of individual/family factors on enrollment and motherhood controlling for state-level variables, cross-lagged effects, and state dependence (the latter presented in Table 7). Two models—without and with state fixed effects—were estimated. Controlling for unobserved state characteristics did not change the estimates of individual/family factors, so we can focus on the results under the model without state FE. Most of our findings about individual/family factors are consistent with the literature and our estimates are more rigorous because we consider school enrollment and

Table 6 Effects of individual/family variables on school enrollment and teen motherhood

Variable	w/o state FE		w/ State FE	
	Enrollment	Teen Motherhood	Enrollment	Teen Motherhood
Age (in months)	–.02**	.02**	–.02**	.02**
	(12.85)	(13.97)	(12.66)	(13.75)
Black	.32**	.13*	.30**	.11*
	(6.93)	(2.40)	(6.23)	(1.96)
Hispanic	.13*	.05	.12*	.03
	(2.43)	(.79)	(2.09)	(.44)
Other race	.16	.00	.08	.05
	(1.46)	(.01)	(.69)	(.38)
Stepfamily	–.07	.15*	–.06	.16*
	(1.16)	(2.09)	(.94)	(2.20)
Single-mother	–.16**	.23**	–.15**	.21**
	(3.21)	(3.87)	(3.03)	(3.56)
Single-father	–.21*	.22^	–.19^	.216^
	(2.19)	(1.93)	(1.95)	(1.88)
Other family	–.28**	.42**	–.27**	.43**
	(5.49)	(6.52)	(5.51)	(6.60)
Income-to-needs	.01	–.04**	.01	–.04**
	(.80)	(3.06)	(.79)	(3.36)
Parental education	.03**	–.03**	.04**	–.03**
	(4.92)	(4.10)	(5.42)	(4.12)
Parental AFDC before 1992	–.05	–.03	–.05	–.03
	(.98)	(.46)	(1.01)	(.42)
Parental AFDC 1992–97	–.08^	.04	–.09*	.07
	(1.80)	(.71)	(2.06)	(1.20)
# siblings	.00	.01	.00	.02
	(.36)	(.98)	(.00)	(1.26)
Urban residence	–.16**	.05	–.11*	.04
	(3.76)	(.94)	(2.47)	(.70)

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text).

Note: Presented are estimates from a model of two interdependent processes (t-ratios are in parentheses). See Appendix for technical notes. The model includes state-level variables, which are presented in Table 7.

** $p < .01$, * $p < .05$, ^ $p < .10$ (two-tail test)

teen motherhood simultaneously. Age is negatively associated with school enrollment (either by graduation or dropout) and positively associated with teen motherhood. Controlling for family structure and SES, blacks are more likely to enroll in school as well as in motherhood. Hispanics are also more likely to be enrolled in school when family background is held constant but they are not more likely to enter motherhood. All non-intact family types increase the probability of entering motherhood and all but stepfamily decrease the probability of being enrolled. While parental education promotes enrollment and depresses motherhood, income-to-needs only depresses motherhood. Controlling for other background variables, parental welfare receipt plays only a small role: its decreasing effect on enrollment is only

Table 7 Effects of state-level variables on school enrollment and teen motherhood

Variable	w/o state FE		w/ state FE	
	Enrollment	Teen Motherhood	Enrollment	Teen Motherhood
Unemployment rate	–.00	–.01	–.06	.02
	(.06)	(.23)	(1.52)	(.49)
Abortion policy index	.03	.07*	–.04	.08
	(1.24)	(2.17)	(.35)	(.71)
Paternity establishment*non-poor	–.05	.03	.06	–.05
	(1.44)	(.64)	(1.02)	(.84)
Paternity establishment*poor	–.12**	.07^	–.02	.00
	(3.65)	(1.68)	(.29)	(.07)
Locator service*non-poor	.02**	–.02*	.06**	–.03
	(2.95)	(2.12)	(3.18)	(1.25)
Locator service*poor	.00	–.01	.04*	–.02
	(.45)	(1.17)	(2.16)	(.95)
Minor mandate index*non-poor	.07^	–.06	.03	–.05
	(1.70)	(1.25)	(.72)	(1.13)
Minor mandate index*poor	.00	–.07	–.02	–.08^
	(.11)	(1.47)	(.35)	(1.66)
Lagged enrollment	1.93**	–.19**	1.92**	–.19**
	(91.66)	(2.90)	(91.80)	(2.96)
Lagged teen motherhood	–.20**	2.13**	–.21**	2.12**
	(2.99)	(59.08)	(3.14)	(57.75)

Data sources: NLSY97 waves 1–4 and retrospective fertility history and various sources of state-level variables (see text)

Note: Presented are estimates from a model of two interdependent processes (t-ratios are in parentheses). See Appendix for technical notes. The model includes individual/family variables, which are presented in Table 6

** $p < .01$, * $p < .05$, ^$p < .10$ (two-tail test)

marginally significant and its effect on motherhood is insignificant. Urban residence decreases enrollment but it does not increase motherhood.

Table 7 presents the results for state-level variables, the cross-lagged effects, and state dependence, controlling for individual/family variables (the effects of which were presented in Table 6). Controlling for state FE did not affect the estimates of unemployment rate and abortion policy. Unemployment rates are not significantly associated with either enrollment or motherhood without or with state FE. Stricter abortion policies increase the probability of motherhood under the model without state FE. The coefficient under the model without state FE is almost of the same magnitude but the standard error is larger, so that it is no longer significant. We cannot rule out the possibility that we might be able to identify the abortion policy effect if we were to have more data.

Turning to the child support and welfare policy variables, the model without state FE appears to produce two types of bias. First, it overestimates the effects of paternity establishment on both enrollment and teen motherhood for the poor and the effects of the minor mandate index on enrollment for the non-poor. The model without state FE finds that an increase in one standard deviation of paternity establishment decreases the log odds of school enrollment by –.12 for girls from poor families. Under the model with state FE, this effect is close to zero, however. The result is similar for teen motherhood. The positive effect of paternity establishment on motherhood under the model without state FE becomes close to zero under the model with state FE. We find the same overestimation bias for minor mandate index for girls from non-poor families. The second type of bias is underestimation. The coefficients for locator service (regardless of poverty status) become larger and significant for enrollment after state FE is controlled. The effect of the interaction between the minor mandate index and poverty becomes marginally significant on teen motherhood after state FE is controlled. Thus, controlling for unobserved state climates, locator service increases all girls' enrollment and minor mandate marginally decreases motherhood for girls from poor families. These results provide some evidence to support our Hypothesis 2 about the depressing effect of welfare policy on motherhood and Hypothesis 3 about the positive effect of child support policy on enrollment.

The test of Hypothesis 4 about the effect of enrollment on motherhood that serves as an indirect route of policy variables to depress teen motherhood is through the specification of the lagged enrollment in the model. The coefficients for lagged enrollment (see the bottom of Table 7) show that lagged enrollment indeed depresses teen motherhood. The coefficient is –.19 and statistically significant under both models without and with state FE. This estimate is identified after controlling for individual/family/state covariates and state dependence of both enrollment and motherhood. Combined with the positive effects of child support policies, the results support our Hypothesis 4 that the higher levels of locator service a state provides keep young girls in school, which depresses teen motherhood. Results in Table 7 also reveal the effects of lagged motherhood on enrollment. Lagged motherhood decreases the probability of enrollment. The coefficient is –.20 without or with state FE. This finding suggests that welfare policy on minor mandates indirectly increases enrollment because minor mandates decrease motherhood marginally, which in turn decreases enrollment.

Finally, the bottom two lines in Table 7 show strong state dependence of both enrollment and motherhood. Note that the estimates of state dependence are obtained after controlling for individual/family/state covariates and cross-lagged feedback of enrollment and motherhood. In turn, the strong significant state dependence shows the importance of its inclusion in the model in order to identify the effects of our key explanatory variables on child support and welfare policies and the reciprocal relationship between enrollment and motherhood.

Summary and discussion

The results we presented above are both confirmatory and new. New findings emerge from our effort to model teenage girls' enrollment and motherhood simultaneously in assessing the impacts of child support enforcement and welfare policies. Little past research has examined the potential reciprocal relationship between teenagers' enrollment and motherhood, taking state dependence and unobserved state climates into account. Our results are also based on very recent data from the experiences of a nationally representative sample of teenagers who went through the sweeping welfare reform from the early 1990s to the early 2000s.

As with all teen fertility data collected by surveys, including births and pregnancies, the NLSY97 data face an underreporting problem. Pregnancies that did not end in live births are usually underreported. If underreporting is selective according to the perceived pressure from social norms, our teen births and motherhood data may overrepresent adolescents who are not conformists. Then our results may be contaminated by such selectivity.

With this caveat, our analysis confirms two sets of previous findings and adds new knowledge. Our results for young women confirm those of Jones et al. (1999) and others in that individual/family factors remain central in determining adolescents' fertility behavior, although we find the strengths and significance levels differ between early and late adolescence. Second, our failure to find evidence for the expected welfare policy effects on nonmarital teenage births confirms the findings of Foster and Hoffman (2001) that welfare policy has little effect on the decisions of teenage girls to have births, after unobserved characteristics of the states are taken into account. It also confirms the findings of Hao and Cherlin (2004) that welfare reform does not reduce teenage pregnancies and births.

Going beyond the past research, our analysis investigates the possibility that welfare policies affect teen motherhood by modeling teen motherhood and school enrollment simultaneously. We extend beyond the event of teen births to the state of motherhood that considers both being pregnant and raising a child. Our model for the interdependence of motherhood and enrollment takes into account state dependence of motherhood, state dependence of enrollment, and the reciprocal relationship between enrollment and motherhood using a cross-lagged specification. Estimates from this model reveal a marginally significant depressing effect on motherhood of minor mandates on attending school and living with parents. Because lagged motherhood reduces current enrollment, welfare policy also indirectly and marginally increases school enrollment, although we fail to identify a direct effect of welfare policy on enrollment.

Although previous aggregate analyses support the depressing effect of child support policies on childbearing of all reproductive-age women (Garfinkel et al. 2003), we fail to find such a direct effect on teenage childbirths and we find an indirect depressing effect of child support policy on teen motherhood through school enrollment. Although our study and the previous research are similar in the economic rationale about the greater costs of children for unwed fathers and in controlling for state fixed effects, they differ in four ways. First, our analysis focuses on teenagers rather than reproductive-age women. Second, it follows individual teenage girls from age 12 to age 19, making use of the monthly trajectories and transitions. Third, we simultaneously model teen motherhood and school enrollment. Last, we include measures of child support policy relevant to teenagers, including paternity establishment and locator service (rather than collection rates and amounts). These measures may convey to teenage girls and young men prospects of strict enforcement of child support if a child is born to them. Our rational for why child support can influence teenage childbirths and motherhood is that young men are more likely to refrain from fathering children under stricter child support policies so that young girls are less distracted from schooling, which in turn reduces the probability of motherhood. We find evidence to support this conjecture. In particular, greater levels of locator service increase school enrollment of teenage girls from both poor and non-poor families and lagged school enrollment reduces the chance of current motherhood.

Child support enforcement imposes economic obligations on biological fathers and conveys the importance of schooling in increasing one's capacity to support a family. Our study suggests that these messages may be reaching young men and women. Welfare policy makers and administrators, understandably, tend to think of child support enforcement as a policy to support low-income children and to recoup some of the costs of providing cash assistance to their mothers. While that is its main focus, child support enforcement may have the ancillary benefit of discouraging young men from fathering children and therefore allowing more young women to remain in school. If so, policy makers may wish to establish outreach programs aimed at teenagers that make the message even clearer. These efforts could potentially reduce nonmarital teen motherhood, which reduces the needs for both welfare support and child support enforcement in the long run.

Technical Appendix

This appendix shows the relationships between multinomial logit parameters based on a cross-classification of two endogenous variables and parameters that fit the transition of each of the two endogenous variables under Yamaguchi's model (1990). The method converts multinomial logit parameters into parameters of interest and their standard errors.

To begin, suppose we have two dichotomous random variables A and B that take the value of either –1 or 1. Cross-classifying the two variables, we have four categories:

		B	
		–1	1
A	–1	P_{00}	P_{01}
	1	P_{10}	P_{11}

For simplicity, we have one covariate, X, and the two lagged endogenous variables with δ's being their parameters. A multinomial logit specification is then:

$$\begin{aligned} log\left(\frac{p_{01}}{p_{00}}\right) &= \delta_{20} + \delta_{21}A_{t-1} + \delta_{22}B_{t-1} + \delta_{23}X \\ log\left(\frac{p_{10}}{p_{00}}\right) &= \delta_{30} + \delta_{31}A_{t-1} + \delta_{32}B_{t-1} + \delta_{33}X \\ log\left(\frac{p_{11}}{p_{00}}\right) &= \delta_{40} + \delta_{41}A_{t-1} + \delta_{42}B_{t-1} + \delta_{43}X. \end{aligned} \tag{1}$$

$A = -1$ & $B = -1$ is the reference category and its parameters, δ_1's, are constrained to 0.

Following Yamaguchi, let μ, ϕ^A, ϕ^B and ϕ^{AB} represent a set of parameters that are defined as the following:

$$\begin{aligned} log\, p_{00} &= \mu - \phi^A - \phi^B + \phi^{AB} \\ log\, p_{01} &= \mu - \phi^A + \phi^B - \phi^{AB} \\ log\, p_{10} &= \mu + \phi^A - \phi^B - \phi^{AB} \\ log\, p_{11} &= \mu + \phi^A + \phi^B + \phi^{AB} \end{aligned} \tag{2}$$

where ϕ^A fits the log-odds of A, ϕ^B fits the log-odds of B, and ϕ^{AB} fits the A and B association. Through algebraic manipulation of equations in (2), we get:

$$\begin{aligned} 4\phi^A &= -\log\left(\frac{P_{01}}{P_{00}}\right) + \log\left(\frac{P_{10}}{P_{00}}\right) + \log\left(\frac{P_{11}}{P_{00}}\right) \\ 4\phi^B &= \log\left(\frac{P_{01}}{P_{00}}\right) - \log\left(\frac{P_{10}}{P_{00}}\right) + \log\left(\frac{P_{11}}{P_{00}}\right) \\ 4\phi^{AB} &= -\log\left(\frac{P_{01}}{P_{00}}\right) - \log\left(\frac{P_{10}}{P_{00}}\right) + \log\left(\frac{P_{11}}{P_{00}}\right) \end{aligned} \tag{3}$$

ϕ^{At}, ϕ^{Bt} and ϕ^{ABt} each is a function of the covariate X and the cross-lagged variables, with λ_s being the corresponding parameters. We multiply each side of the equations by 4:

$$\begin{aligned} 4\phi^{A_t} &= 4\lambda_0^A + 4\lambda_1^A\, A_{t-1} + 4\lambda_2^A\, B_{t-1} + 4\lambda_3^A X \\ 4\phi^{B_t} &= 4\lambda_0^B + 4\lambda_1^B\, A_{t-1} + 4\lambda_2^B\, B_{t-1} + 4\lambda_3^B X \\ 4\phi^{AB_t} &= 4\lambda_0^{AB} + 4\lambda_1^{AB}\, A_{t-1} + 4\lambda_2^{AB}\, B_{t-1} + 4\lambda_3^{AB} X \end{aligned} \tag{4}$$

where λ_1^A is the first-order state-dependence for A_t, λ_2^A is the lagged causal effect of B_{t-1} on A_t, λ_3^A is the parameter related to the covariate. Similarly, λ_1^B is the lagged causal effect of A_{t-1} on B_t, λ_2^B is the first-order state-dependence for B_t, λ_3^B is the parameter related to the covariates. λ^{AB}'s are parameters for the association between A_t and B_t.

Substituting Eqs. (1) and (4) into Eq. (3), we can solve for the parameters corresponding to the covariate. For example, for the parameters corresponding to A_{t-1}, we have Eq. (5) where the parameters that fit A, B, and their association AB can be expressed as linear combinations of the multinomial logit parameters:

$$\begin{aligned} 4\lambda_1^A &= -\delta_{21} + \delta_{31} + \delta_{41} \\ 4\lambda_1^B &= \delta_{21} - \delta_{31} + \delta_{41} \\ 4\lambda_1^{AB} &= -\delta_{21} - \delta_{31} + \delta_{41} \end{aligned} \tag{5}$$

Since the variance of a linear combination of random variables is the sum of the variance of each variable and twice the covariance among them (adjusting for the scaling and the signs), the corresponding variance for a particular parameter can be expressed as linear combinations of elements in the variance-covariance matrix of the multinomial logit estimation. For the variances of parameters corresponding to A_{t-1}, we get:

$$\begin{aligned} Var(4\lambda_1^A) &= Var(\delta_{21}) + Var(\delta_{31}) + Var(\delta_{41}) - 2Cov(\delta_{21}, \delta_{31}) - 2Cov(\delta_{21}, \delta_{41}) \\ &\quad + 2Cov(\delta_{31}, \delta_{41}) \\ Var(4\lambda_1^B) &= Var(\delta_{21}) + Var(\delta_{31}) + Var(\delta_{41}) - 2Cov(\delta_{21}, \delta_{31}) + 2Cov(\delta_{21}, \delta_{41}) \\ &\quad - 2Cov(\delta_{31}, \delta_{41}) \\ Var(4\lambda_1^{AB}) &= Var(\delta_{21}) + Var(\delta_{31}) + Var(\delta_{41}) + 2Cov(\delta_{21}, \delta_{31}) - 2Cov(\delta_{21}, \delta_{41}) \\ &\quad - 2Cov(\delta_{31}, \delta_{41}) \end{aligned} \tag{6}$$

We can obtain the variance of other parameters in the same manner.

References

Astone, N. M., & Upchurch, D. M. (1994). Forming a family, leaving school early, and earning a GED: A racial and cohort comparison. *Journal of Marriage and the Family 56*, 759–771.

Blank, R. M. (2002). Evaluating welfare reform in the United States. *The Journal of Economic Literature 15*, 1105–1166.

Bloom, D. E., Conrad, C., & Miller, C. (1998). Child support and father's remarriage and fertility. In I. Garfinkel, S. S. McLanahan, D. R. Meyer, & J. A. Seltzer (Eds.), *Fathers under fire: The revolution in child support enforcement*. (pp. 128–156). New York: Russell Sage.

Fein, D. J. (1999). Will welfare reform influence marriage and fertility? Early evidence from the ABC demonstration. Bethesda MD: Abt Assoc.

Finkel, S. F. (1995). *Causal analysis with panel data*. Thousand Oaks CA: Sage Publications.

Foster, E. M., & Hoffman, S. D. (2001). The young and the not quite so young: Age variation in the impact of AFDC benefits on nonmarital childbearing. In L. L. Wu, & B. Wolfe (Eds.), *Out of wedlock: Causes and consequences of nonmarital fertility*. (pp. 173–201). New York: Russell Sage.

Freeman, R. B., & Waldfogel, J. (1998). Does child support enforcement policy affect male labor supply?. In I. Garfinkel, S. S. McLanahan, D. R. Meyer, & J. A. Seltzer (Eds.), *Fathers under fire: The revolution in child support enforcement*. (pp. 94–127). New York: Russell Sage.

Garfinkel, I., Huang, C.-C., McLanahan, S., & Gaylin, D. (2003). Will child support enforcement reduce nonmarital childbearing? *Journal of Population Economics 16*(1), 55–70.
Garfinkel I., McLanahan S. S., Meyer D. R., & Seltzer J. A. (Eds.). (1998). *Fathers under fire: The revolution in child support enforcement*. New York: Russell Sage.
Gottschalk, P. (1990). AFDC participation across generations. *American Economic Association Papers and Proceedings 80*(2), 367–371.
Gottschalk, P. (1992). The intergenerational transmission of welfare participation: Facts and possible causes. *Journal of Policy Analysis and Management, 11*(2), 254–272.
Hao, L. (1997). Using a multinomial logit specification to model two interdependent processes with an empirical application. *Sociological Methods and Research 26*(1), 80–117.
Hao, L., Astone, N. M., & Cherlin, A. J. (2004). Adolescents' school enrollment and employment: Effect of state welfare policies. *Journal of Policy Analysis and Management 23*, 697–721.
Hao, L., & Cherlin, A. J. (2004). Welfare reform and teenage pregnancy, childbirth, and school dropout. *Journal of Marriage and the Family 66*, 179–194.
Hardy, J. B., Astone, N. M., Brooks-Gunn, J., Shapiro, S., & Miller, T. L. (1998). Like mother, like child: Intergenerational patterns of age at first birth and associations with childhood and adolescent characteristics and adult outcomes in the second generation. *Developmental Psychology 34*(6), 1220–1232.
Haveman, R., & Wolfe, B. (1994). Succeeding generations: On the effects of investments in children. New York: Russell Sage.
Hsiao, C. (2003). *Analysis of panel data* (2nd ed.). New York: Cambridge University Press.
Huang, C.-C. (2001). The impact of child support enforcement on nonmarital and marital births: Does it differ by racial and age groups? *Social Service Review 76*(2), 275–301.
Huber, P. J. (1967). The behavior of maximum likelihood estimate under non-standard conditions. In *Proceedings of the fifth Berkeley symposium on mathematical statistics and probability 1*. (pp. 221–233). Berkeley CA: University of California Press.
Jones, A. S., Astone, N. M., Keyl, P. M., Kim, Y. J., & Alexander, C. S. (1999). Teen childbearing and educational attainment: A comparison of methods. *Journal of Family and Economic Issues 20*(4), 387–417.
Joyce, T., & Kaestner, R. (1996). The effect of expansions in Medicaid income eligibility on abortion. *Demography 33*(2), 181–192.
Joyce, T., Kaestner, R., & Kwan, F. (1998). Is Medicaid pronatalist? The effect of eligibility expansions on abortions and births. *Family Planning Perspectives 30*(3), 108–113.
Kaestner, R., Korenman, S., & O'Neill, J. (2003). Has welfare reform changed teenage behavior? *Journal of Policy Analysis and Management 22*, 225–248.
Meyer, D. R. (1998). The effect of child support on the economic status of nonresident fathers. In I. Garfinkel, S. S. McLanahan, D. R. Meyer, & J. A. Seltzer (Eds.) *Fathers under fire: The revolution in child support enforcement*. (pp. 67–93). New York: Russell Sage.
Moffitt, R. A. (1998). The effect of welfare on marriage and fertility. In R. A. Moffitt (Ed.) *Welfare, the family and reproductive behavior*. (pp. 59–97). Washington, DC: National Academies Press.
Offner, P. (2003). *Teenagers and welfare reform*. Washington, DC: Urban Institute.
Plotnick, R., Garfinkel, I., McLanahan, S., & Ku, I. (2005). The impact of child support enforcement policy on nonmarital childbearing. Working paper #2006–2009. Retrieved at http://www.evans.washington.edu/webtools/working_papers/Evans-Faculty-Paper.php?id=44.
Quint, J. C., Bos, J. M., & Polit, D. F. (1997). *New chance: Final report on a comprehensive program for young mothers in poverty and their children*. New York: Manpower Demonstration Research Corporation.
Wu, L. L. (1996). Effects of family instability, income and income insecurity on the risk of a premarital birth. *American Sociological Review 61*(3), 386–406.
Wu, L. L., & Martinson, B. C. (1993). Family structure and the risk of a premarital birth. *American Sociological Review 58*(2), 210–232.
Yamaguchi, K. (1990). Logit and multinomial logit models for discrete-time event-history analysis: A causal analysis of interdependent discrete-state processes. *Quality & Quantity 24*, 323–341.