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Legislating Love: The Effect of Child Support and Welfare Policies on Father–child Contact*

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Abstract. Reducing non-marital childbearing and making nonresidential fathers take greater responsibility for their children were identified as goals of numerous policy changes since the 1980s. Child-support award rates for children born to unmarried parents have been quite low historically, leading lawmakers to focus on increasing both award and payment rates for this group. Nonmarital fathers are also much less likely to have contact with their children. Although evidence suggests that policy efforts increase child support awards and receipt, the link between child support policies, child support outcomes, and father-child contact has received less attention. This paper uses data from the Survey of Income and Program Participation on children born between 1985–1997 to investigate the relationship between child-support award and receipt and the amount of contact that fathers have with their non-residential children. Since it is likely that both of these behaviors are, in part, determined by unobservable characteristics of the father, we estimate an instrumental variables Tobit model. The model is identified by our assumption that child support policy variables can impact child support awards and payments, but father-child contact cannot be directly legislated. Our results suggest that there are unintended, but desirable effects of child support establishment and collection. Policies to collect child support not only increase financial resources to families, but through their impact on payments increase visitation and contact between these children and their fathers. The estimated impact of receiving child support on contact is more than 27 days per year.

Keywords: child support, father-child contact, nonmarital fathers, IV Tobit

JEL Classification: J13

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1. Introduction

As non-marital childbearing increased dramatically over the last two decades—from 18% of all births in 1980 to almost one-third of births in 1998 (National Center for Health Statistics, 1999)—policy makers focused their attention on ways to reduce non-marital childbearing and make absent fathers take greater responsibility for their children. Indeed, these goals explicitly motivated numerous welfare and child support policies adopted during the 1980s and 1990s. Child-support award rates for children born to unmarried parents have been quite low historically. Nonmarital fathers are also much less likely to have contact with their children (Laura Argys and Elizabeth Peters, 2001). Although evidence suggests that policies can increase child support awards and receipt (Andrea Beller and John Graham, 1993; Irwin Garfinkel and Philip Robbins, 1994; Richard Freeman and Jane Waldfogel, 2000; Argys, Peters and Donald Waldman, 2001), the link between child support policies, child support outcomes, and father–child contact has received less attention (see, however, recent work by Judith Seltzer, McLanahan and Hanson, 1998; Argys and Peters, 2001; Elaine Sorensen and Kate Pomper, 2002; and Garfinkel et al., 2002).

This paper investigates the relationship between child-support and the amount of contact non-marital children have with their fathers. Since it is possible that both of these behaviors are, in part, determined by unobservable characteristics of the father, we estimate an instrumental variables Tobit model to account for possible endogeneity. The joint estimation of child support and father–child contact allows us to answer the questions: does the enforcement of financial support awards have a secondary effect of increasing father–child contact for children whose parents never married? Can we, in effect, legislate love? Section 2 discusses the expected relationship between child support and time involvement based on several different theoretical perspectives. The data are described in Section 3. Section 4 briefly outlines changes in government policy regarding non-marital fertility and child-support enforcement. Section 5 explains the econometric methodology. The results are described in Section 6, and Section 7 concludes.

2. Modeling the relationship between policies, money and time involvement

Recent federal and state legislation have authorized or required more aggressive measures to establish and enforce child support awards. While it is straight-forward to adopt policies to collect money, emotional attachment cannot be legislated. Thus, any effect of policies on fathers' non-pecuniary involvement with their children must occur indirectly. In our analyses we examine one specific type of non-pecuniary involvement: the amount of contact between non-marital children and their non-residential fathers. We investigate whether policies designed to improve child support collections will also increase this contact.

A number of theories attempt to explain how forcing fathers to take more financial responsibility could lead them to provide more non-pecuniary support as well. For example, psychological theories about the "breadwinner role" argue that the perceived success in the role of financial supporter may increase fathers' confidence in taking on non-financial paternal roles as well. Similarly, when mothers act as "gate-keepers", child support payments by the father may be required to gain access to his child. In this case, contact may carry a high price. If, however, the father is forced to pay child support by the state, the marginal cost of seeing his child falls, and father–child contact should increase. In addition, there is the idea that establishing paternity and legislating financial responsibility will alter the expectations of the father early on in the relationship, and that could have longer terms effects on fathers' behaviors.

The discussion above suggests ways in which policies can play a causal role in affecting financial as well as non-financial father involvement. However, it is important to note that correlation between financial and non-financial involvement does not necessarily imply causation. For example, Yoram Weiss and Robert Willis (1985) and Argys and Peters (2003) argue that contact may have a causal effect on child support payments, because contact allows fathers to monitor how the mother is spending that child support. In addition, it is easy to imagine unobserved characteristics of fathers that would increase both the willingness to provide financial support and the desire to be emotionally involved in the child's life. Fortunately, the existence of different child support policies that are likely to affect child support, but not to affect contact directly can help identify a causal relationship between child support awards and payments and father–child contact.

A logical question that follows from the discussion above is "what is the value of estimating a structural model (with all of its inherent complications in identifying a causal relationship), if what we are interested in is the effect of child support policies on contact?" The answer to that general question has been outlined by Robert Lucas (1981), James Heckman, Robert Lalonde, and Jeffrey Smith (1999), and others. Specifically, what we get from a reduced form analysis is the impact of *particular* child support policies on outcomes such as father–child contact. However, without understanding the pathway through which these policies affect contact, we are not able to assess the impact of other proposed child support policies (that may not yet be in effect) on father–child contact. In this paper we propose that the pathway is through the policies' impact on child support payments, which then, in turn, motivate fathers to have more contact. The structural model gives an estimate of the effect of an additional dollar of child support on father–child contact. This estimate can then be used to compare the effectiveness of different child support policies on contact, if the impact of the policies on child support payments and awards is the only thing that is known.

3. Data description

Data about paternal involvement with children come from the 1992, 1993, and 1996 panels of the Survey of Income and Program Participation (SIPP). Each panel of SIPP includes a nationally representative sample ranging from 20,000 to 37,000 households. The households are interviewed every 4 months. In the 1992 and 1993 panels, interviews continue over period of two and a half years each. For the 1996 panel, interviews continue for 4 years. Core information asked at each interview (or wave) includes household composition, basic demographics, income, labor force participation, and receipt of welfare and other public transfers. Each wave also includes additional modules that collect detailed information about specific topics.[1] The module topics differ across waves.

Most of the data used in this paper come from the child support topical modules (wave 9 for the 1992 and 1993 panels, wave 5 for the 1996 panel). These modules identify children with a non-resident father and ask detailed questions about paternal involvement. Specifically the survey asks whether the child was covered by a child support agreement, whether any child support had been received in the past year, and how much personal contact the child had with the father in the past year. The survey also asks about other forms of financial assistance, such as health insurance coverage, that are not used in this analysis.

The sample used in this analysis is limited to children (a) whose biological parents never married each other, (b) whose biological father was alive, and (c) who were living with their biological mother at the time of the child support module interview. Information about the first criteria was determined by comparing the date of birth of the children to information given in the marital history module about the mother's marriage and divorce dates. The state in which the child resided at birth was found by comparing the date of birth to the mother's migration module. The sample used in this paper includes 2189 children born between 1985 and 1997, a period of major change in child support policies.[2]

Due to the design of the survey, data on father–child contact was not collected for all children in a family. Specifically, when there were multiple children in a household, with different amounts of contact, information about father–child contact was generally only collected for one child.[3] Hence contact data are missing for about 30 percent of the children in our initial sample. However, the number of *families* for which no contact information is provided is smaller.[4] We use data for multiple children, when available to increase the efficiency of the estimation. Our estimates will be biased if fathers care differently about their oldest child relative to their younger children, but neither theory nor prior empirical analyses shed light on the direction of the bias.[5]

Characteristics of our sample of children who were born to unmarried parents between 1985 and 1997 are reported in the first column of Table 1. Most of the variables are measured during the child support module. Just under 40 percent of the sample is composed of African–American children, and 16 percent of the sample is Hispanic. There were slightly fewer than two non-marital children residing in the average household. The age of the sample children ranges from 0 to 12 with a mean near 5 years of age. A little more than one quarter of the children were born to teenage mothers, and, on average, their mothers had less than a high school education. About 12 percent of mothers were married at the time of the child support module (not to the father of her child). The data include a measure of the mother's

Table 1. Unweighted sample means by child support status.

Description of variable	Full sample	Child support not received	Received some child support
Individual and family characteristics			
Age of child (years)	4.84(3.22)	4.63(3.26)	5.38(3.06)
Male child	0.497	0.492	0.510
Hispanic (mother)	0.164	0.189	0.097
Black, non-hispanic (mother)	0.386	0.417	0.304
Mother's education (highest grade completed)	11.77(2.06)	11.62(2.15)	12.16(1.74)
Number of nonmarital children in household	1.85(1.00)	1.92(1.05)	1.67(0.82)
Mother currently married	0.116	0.099	0.162
Mother teen at time of birth	0.280	0.282	0.274
Father lives in the same state as child	0.470	0.354	0.771
Father lives in the same city/county as child	0.334	0.261	0.526
Father's residence unknown	0.233	0.302	0.053
Mother born outside the USA	0.088	0.102	0.053
Mother's birthplace unknown	0.156	0.160	0.147
Father–child interactions			
Any child support agreement	0.352	0.159	0.856
Any child support received	0.277	0.000	1.000
Any father–child contact	0.360	0.237	0.682
Days of father–child contact in past year	27.47(70.44)	20.33(65.31)	46.14(79.43)
Sample size	2189	1583	606

Standard errors for continuous variables are in parentheses.

place of birth, and we identify children whose mothers were born outside of the U.S. Nearly 9 percent of the children in our sample have foreign-born mothers. Finally, we include measures of the father's place of residence relative to the child's. The father's residence is known for about three quarters of our sample. Forty-seven percent of the children in our sample live in the same state as their father, and 33 percent are known to live in the same city.

The lower panel of Table 1 reports the means for child support measures and father–child contact. Thirty-five percent of the children in our sample were covered by some sort of child-support agreement. Independent of whether any such agreement existed, the module also asked whether financial support had been received from the father in the past year. More than one-quarter of all children in the sample received some cash transfer in the past year. Predictably, children whose fathers paid child support are more likely to have been covered by a child support award, though nearly 15 percent of children receiving child support do not have an award.

We also report measures of father–child contact. The interviewer asked how much time had been "spent visiting the other parent" in the past year. Answers range from no contact at all, to contact nearly every day. About one-third of the children in our sample saw their fathers at least once during the year prior to the interview. The

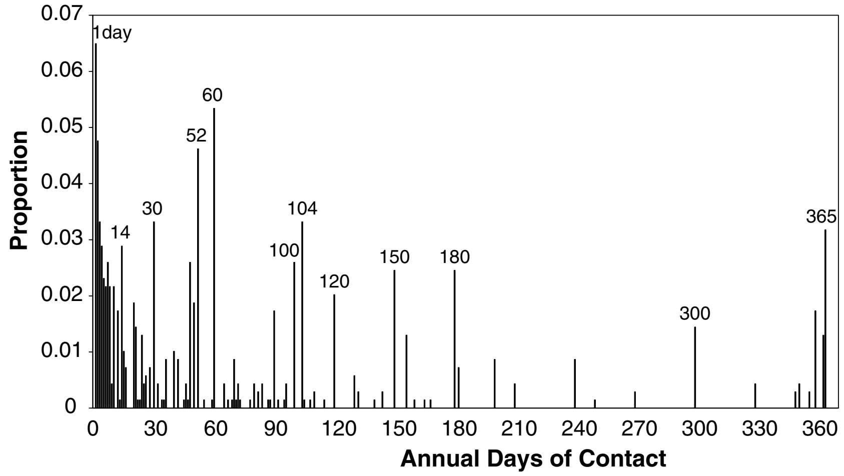

Figure 1. Frequency of annual father–child contact among children with any contact for children born to unmarried parents.

average number of days of annual contact was 27. Fathers who are not paying child support are much less likely to have contact with their children (20 vs. 46 days per year).

The distribution of the number of annual days of father–child contact for those children in our sample who had any contact with their father in the year prior to the survey is depicted in Figure 1. Among children with contact, 6.5 percent saw their father only one day in the past year and more than 20 percent saw their fathers less than 7 days per year. At the other extreme, over three percent saw their father every day. Figure 1 also demonstrates that reports of the amount of father–child contact is subject to a problem also found in reports of hours of work, namely that answers are likely to be rounded. For example 300 days of contact is much more likely to be the answer than 299 days. This type of clustering of responses is evident in Figure 1 with high frequencies of responses at 52 days (once a week) and at 60 days (5 days per month), and at other benchmark amounts.

To address the possible endogeneity of child support award and receipt status, our model incorporates child support and welfare policies that are likely to affect child support, but not to affect contact directly. To link the child-level data with the state-level policy variables (described in the next section) we use migration history data to identify the state of residence at the time of the child's birth and at the time of the child support interview.

4. Government policies aimed at child support collection

Historically, the primary goal of child support policies was to ensure that children with a living but absent parent would receive adequate financial support from that parent (primarily the father). Although most divorced fathers pay some child support, fathers with children born outside of marriage are much less likely to provide support. In 1995, 75 percent of divorced mothers, but only 44 percent of never-married mothers were owed child support (Timothy Grall, 2002). There are a number of reasons for these low rates of support. First, fathers with non-marital children are more likely to have low levels of education and job skills and, thus, low earnings capacity. Second, compared to divorced fathers, non-marital fathers are less likely to have ever lived with their children, and as a consequence may have lower levels of attachment to their children. Third, the legal requirement to establish paternity before child support can be awarded increases the costs of trying to obtain child support from non-marital fathers.

Under Title IV-D of the Social Security Act, states are required to provide child support services such as facilitating paternity establishment, locating absent parents, establishing child-support orders, and enforcing and modifying child-support orders. To indicate a state's relative effectiveness in these roles we use two state-level measures. First, we calculate the average amount of administrative expenditure by the Child Support Enforcement (CSE) office made on behalf of each family in the CSE caseload. Table 2 reports the mean values of the policy variables in the year of birth of each child in our sample (between 1985 and 1997) as well as at the time of the child support interviews (between 1994 and 1997). As measured in the year of birth, states were spending an average of \$162 per CSE case (measured in constant 1995 dollars). Over this time period the median state CSE expenditure level was slowly rising. The lowest expenditure was \$30 per case and the highest was over \$900.

To measure the effectiveness of these expenditures we include the ratio of the amount of child-support collections by the CSE office to the amount of administrative expenditures. This measure of agency effort and effectiveness varies substantially across states and over time. Between 1985 and 1997, all states at least broke-even, and even the worst states were improving over this period. The poorest performing state in our sample collected only \$133 for every \$100 in administrative expenditures. The most efficient states collected over \$1,000 for every \$100 expended. (For a more detailed description of these two variables, see Argys and Peters, 2003).

In response to a perception that child-support award amounts were not uniform and were often inadequate, states were required by federal legislation to implement presumptive child support guidelines. Because the adoption of guidelines has been found to increase the likelihood and amount of child-support awards among unmarried parents (Argys, Peters and Waldman, 2001), we include a dichotomous measure indicating whether each state had implemented presumptive child support guidelines. Although these federal regulations were imposed on all states, the ways in which the regulations are implemented and the timing of implementation varied across states. The majority of states adopted such guidelines between 1985 and 1990 (for more detail see Argys, Peters and Waldman, 2001).[6]

Table 2. Definitions and means of state-level variables.

Variable	Definition	Mean at time of birth	Mean at time of interview
CSE administrative expenditures	The amount of annual child support enforcement (CSE) administrative expenditures per case (1995 dollars).	162.21 (74.13)	238.54 (132.19)
CSE collections to expenditures	The ratio of annual child support dollars collected by the CSE agency to annual administrative expenditures.	3.71 (1.69)	3.88 (1.59)
Child support guidelines	A dichotomous variable equal to 1 if the state had adopted statewide advisory or presumptive child support guidelines.	0.859	–
UIFSA	A dichotomous variable equal to one if the state had enacted a statute facilitating child support enforcement across state lines.	0.139	0.646
Genetic testing laws	A dichotomous variable equal to one if the state had enacted laws requiring genetic paternity testing at the request of the principal.	0.891	–
Paternity ratio	The number of paternities established by the CSE IV-D agency divided by the number of out-of-wedlock births.	0.404 (0.210)	–
AFDC benefits	Maximum monthly amount or AFDC provided by the state to a family of three (1995 dollars).	444.36 (172.33)	381.40 (149.30)

Note: Standard deviations for continuous variables are in parentheses.

Another major development in child support award establishment and enforcement has been inter-state cooperation. Historically, when a father moved out of state, the mother had to initiate a new child-support request in his new state of residence. Sometimes, multiple child support awards could accrue for the same father–child pair. States began addressing this issue with several cooperative agreements that facilitated the enforcement of child support across state lines. The most recent and wide sweeping of these is the Uniform Interstate Family Support Act (UIFSA). Among states adopting UIFSA, a child support award decided in one state is enforceable in all other states. The first states adopted UIFSA in 1992, and other states followed in subsequent years. Only 14 states had not adopted UIFSA by 1997.[7]

When parents are unmarried at the time of the child's birth, paternity must be established before a formal child support agreement can be imposed. We measure the overall effectiveness of state paternity establishment efforts as the rate of paternity establishment for the state (number of paternity cases established per out of wedlock birth) in each year.[8] Nationally the paternity rate has been steadily increasing between 1985 and 1997. In our sample, states legally established 40 paternities for each 100 births to unmarried mothers on average. We also define a dichotomous variable equal to one if the state had enacted laws allowing the results of genetic tests to be used to legally determine paternity. The introduction of genetic testing in most states occurred during the 1980s.

One of the goals of increasing child support is to reduce the cost of public transfers to children with a non-resident parent, and explicit provisions now link child support and welfare benefits. The law requires all public assistance recipients to cooperate with the child support agency in establishing paternity and child support orders and enforcing child support. Non-cooperation is grounds for sanctions in the form of reduced welfare benefits. Theoretically, the cooperation requirement should increase child support for welfare recipients, but the severity and likelihood of imposing these sanctions varies across states and time. Also, welfare generosity may indirectly affect the probability of receiving child support, as lower benefit levels may lead mothers to put more pressure on fathers for support. We use the maximum real monthly AFDC benefits for a family of three to represent welfare generosity.

In our empirical model, these policies can be used as instruments to predict the probability of a child support award or the receipt of child support payments and to test for endogeneity. There are two requirements for a good instrument. First, the variable must be correlated with child support awards and receipt, and second the variable should not directly affect father–child contact. We believe that a strong case can be made that CSE expenditures and collections, guidelines, and interstate cooperation only have a direct effect on child support. However, it may be possible that paternity establishment policies might also directly affect the propensity for a father to want to interact with his child. For example, some fathers may be more likely to want to interact with their child when paternity is assured through genetic testing. Similarly, if welfare policies affect residence decisions of fathers, these policies might affect contact as well as child support. Fortunately, this issue can be investigated empirically. By maintaining the assumption that at least two policy variables meet the criteria for being a proper instrument (e.g., CSE expenditures and collections), other potential instruments can be tested to see if they have any direct effect on contact.[9]

In our empirical work we link these policies to the child's data by their state of residence both at the time of the birth of the child and at the time of the interview in which child support and contact data are collected. Since the probability of a child support award for a child born to unmarried parents is highest in the first year after the child is born (Argys, Peters and Waldman, 2001) in the child support award equation we measure all policies in the year of the child's birth. In the child support receipt equation, we include measures of child support collection policies (CSE Administrative Expenditures, CSE Collections to Expenditures, and UIFSA) and welfare benefits measured at the time of the child support module interview. Other policies related to the establishment of paternity and awards (Child Support Guidelines, Genetic Testing Laws and the Paternity Ratio) are measured in the year of the child's birth in the child support receipt equation.

5. Methodology

We begin our examination of the effect of child support and welfare policies on father–child contact by estimating reduced-form equations. Our data contain a continuous measure of the number of days of father–child contact reported in the past year. Because a substantial fraction of the children in our sample have no contact with their fathers, the appropriate model is a Tobit model of father–child contact. The Tobit model assumes that there is a latent continuous normally distributed variable which is not observed but which gives rise to an observed value above or below a limit:

$$y_t^* = \underline{x}_t'\underline{\beta} + \varepsilon_t, \quad y_t = \max(0, y_t^*), \quad \varepsilon_t \sim N(0, \sigma^2). \tag{1}$$

The subscript t denotes a child, y^* is the latent variable representing the father's desire to spend time with his child, y is observed amount of father–child contact. The explanatory variables are $\underline{x}$, the disturbance is ε, the structural coefficients are $\underline{\beta}$, showing the effect of the explanatory variables on the latent outcome, and σ is the standard deviation of the disturbance. In the reduced form, the vector x includes both family and child characteristics and the policies of interest.

To understand the route through which policies affecting child support may affect father–child contact, we then jointly estimate the relationship between child support award and receipt and contact between a father and his child. This is a Tobit model with two potentially endogenous variables and is estimated by Generalized Method of Moments (GMM) using instrumental variables (IVs).[10]

The first set of expectations that can be used to define orthogonality conditions (o.c.) for GMM (Lars Hansen, 1982) is based on the expectation of the observed dependent variable, y.

$$E(y_t) = \underline{x}_t'\underline{\beta}\Phi(\underline{x}_t'\underline{\beta}/\sigma) + \sigma\varphi(\underline{x}_t'\underline{\beta}/\sigma). \tag{2}$$

The IVs include all exogenous variables in the Tobit model of father–child contact and the additional policy variable that are used to estimate the endogenous right-side variables—child support award and child support receipt. As in two-stage least squares, it is not necessary to estimate those equations explicitly to use the IVs. The vector of all acceptable exogenous IVs is $\underline{z}_t$. The first o.c. uses the expectation of y:

$$E[y_t - \sigma\{(\underline{x}_t'\underline{\beta}/\sigma)\Phi(\underline{x}_t'\underline{\beta}/\sigma) - \varphi(\underline{x}_t'\underline{\beta}/\sigma)\}] = \underline{0}. \tag{3}$$

Define a dummy variable d_t which is 1 if $y^* > 0$ and 0 if $y^* \leq 0$. This is a probit dependent variable. Its expectation is

$$E(d_t) = \Phi(\underline{x}_t'\underline{\beta}/\sigma) \tag{4}$$

from which

$$E(\underline{z}_t(d_t - \Phi(\underline{x}_t'\underline{\beta}/\sigma))) = \underline{0}. \tag{5}$$

The o.c. as written are valid, but not the best for estimation, because the parameter σ is in the denominator, making the numerical estimation more difficult. Therefore, define parameters $\delta = \beta/\sigma$ and substitute into the equations. The statistical procedure estimates δ and σ then transforms back to β and σ using the delta method (Greene, 2000, p. 118). The o.c. are

$$E[y_t - \sigma\{(\underline{x}_t'\underline{\delta})\Phi(\underline{x}_t'\underline{\delta}) - \varphi(\underline{x}_t'\underline{\delta})\}] = \underline{0}. \tag{6}$$

$$E(\underline{z}_t(d_t - \Phi(\underline{x}_t'\underline{\delta}))) = \underline{0}. \tag{7}$$

Note that no parameters are in the denominator, and σ in particular is isolated in one equation. This form of the estimation is particularly convenient for computation.

In the present application, there are three complications: (a) there is a potentially endogenous variable among the explanatory variables, (b) there are more than one instrumental variables with which to estimate the model and (c) the observed dependent variable might be heteroscedastic for reasons of the method of measuring it (in addition to its inherent structural heterogeneity).

The potentially endogenous variables are whether there is any child support agreement and whether any child support was received in the past year. These may be endogenous to contact with the father either because unmeasured reasons for the existence of a child support agreement or payment could increase contact or because unmeasured reasons for contact could alter the probability of a child support agreement or payment.

There are several IVs for the existence of a child support agreement and child support payment. These IVs are the policies and state-level indicators described in Section 2: AFDC generosity, CSE administrative expenditures per case, the ratio of CSE collections to expenditures, the paternity ratio, guidelines, UIFSA and genetic testing provisions that will have a direct effect on whether or not the child support award is determined and enforced. However, none of these variables are expected to affect father–child contact directly. The GMM setup implies an optimal weighting matrix to combine all of this information and a test statistic for the null hypothesis that all extra IVs beyond those needed for exact identification are valid. (See Hansen, 1982 and Russell Davidson and James MacKinnon, 1993, pp. 614–617.) The extra IVs allow for a non-specific test of the structure and the validity of the IVs in addition to allowing for more efficient estimation of the parameters under the null hypothesis that all IVs are valid.

Recall from Figure 1 the clustering of reports of father–child contact. This implies that there is heteroscedasticity. It is quite difficult to deal with this heteroscedasticity when estimating the Tobit model using Maximum Likelihood Estimation in which a complete stochastic specification is needed. Under the assumption that the measurement errors in the dependent variable are zero-mean and unrelated to the explanatory variables (there is no variable that would predict the tendency to round an answer given the true number of days of contact), the expectations are still valid and are all that are needed to specify GMM. Further, GMM automatically controls for heteroscedasticity (see Greene, 2000; Davidson and MacKinnon, 1993; Hansen, 1982). In the present application, this is a major advantage. Using the o.c. in equations (5–6) and the theory of GMM, we estimate the parameters and test the overidentifying restrictions.

6. Results

6.1. Reduced-form model

Coefficients from two reduced-form models are reported in Table 3. In the discussions that follow, we convert these coefficients into marginal probabilities on observed days of contact.[11] Our results indicate that neither characteristics of the child nor the parents are good predictors of the amount of annual contact between the father and his child. Specifically, there are no significant differences in the number of days of contact for children of different ages, gender, or race. Consistent with findings from other work (Argys and Peters, 2001) father–child contact is also unaffected by many of the characteristics of the parents and family such as the place of mother's birth and the current marital status and education of the mother. There are, however, some characteristics of the family that do matter. For instance, children of Hispanics have lower levels of contact as compared to children of white mothers, and father–child contact decreases as the number of nonmarital children in a family increases. As expected, father–child contact also differs significantly by

Table 3. Reduced-form tobit regressions of annual father–child contact for children born to unmarried parents.

Variable	Annual days of father–child contact Tobit coefficients: Child support policies measured in the year of the child's birth	Child support policies measured at the time of child support payment
Age of child (years)	0.10 (2.62)	−1.49 (3.49)
Male child	−5.39 (5.61)	−5.78 (5.59)
Hispanic (mother)	−19.78** (9.39)	−16.43* (9.36)
Black, non-hispanic (mother)	−10.60 (6.86)	−9.28 (6.95)
Mother's education (highest grade completed)	1.80 (1.48)	1.68 (1.48)
Number of nonmarital children in household	−8.66*** (3.45)	−9.13*** (3.16)
Mother currently married	−1.39 (9.43)	−1.63 (9.49)
Mother teen at time of birth	−5.46 (6.59)	−5.92 (6.57)
Father lives in the same state as child	93.24*** (9.76)	90.67*** (9.71)
Father lives in the same city/county as child	25.47*** (8.90)	25.61*** (8.90)
Mother born outside the USA	−0.11 (13.25)	0.43 (13.21)
Birth year trend	−1.82 (2.74)	−2.91 (3.53)
Genetic testing laws	4.77 (12.17)	6.03 (12.16)
Paternity ratio	7.40 (15.99)	7.04 (15.37)
Child support guidelines	4.18 (11.55)	6.31 (11.39)
AFDC benefits[a]	6.14*** (2.14)	6.00*** (2.09)
CSE administrative expenditures[a]	−2.48 (5.17)	0.85 (2.73)
CSE collections to expenditures[a]	3.14* (1.88)	5.91*** (2.02)
UIFSA[a]	22.76** (10.88)	7.85 (9.88)
Sample size	2189	2189
Chi-squared statistic (7 d.f.)	19.547***	23.759***

Notes: Reported Tobit coefficients are for the latent variable. Standard errors of the Tobit coefficients adjusted for clustering at the household level are reported in parentheses. Regression also includes a constant and indicators for missing mother's place of birth and missing father's residence. Chi-square statistic is for a test of the joint significance of the policy variables.

[a]Measured in the year of the child's birth in column one and in the year of child support collection in column two.

*: $p < 10$, **: $p < .05$, ***: $p < .01$.

proximity. A child who lives in the same state as his father will have 33 days more contact per year than a father and child who reside in different states. If the father and child also live in the same city, contact increases by an additional 10 days.

In column one, all of the policy variables are measured at the time of the child's birth, whereas in column two, the variables associated with child support payments (welfare benefits, child support expenditures and collection and UIFSA) are measured at the time that contact and child support payments are reported.[12] Estimates from the reduced-form models in both columns show that child support and welfare policies do have an impact on father–child contact. In each of the models, the policy variables are jointly significant at the one percent level. The coefficients on some of the policies are more precisely estimated than are others. For example, contact is increased as states raise their welfare benefits and improve their collection of child support. In addition, more father–child contact takes place among children who were born in states that had already implemented the UIFSA. Children born in states after UIFSA was adopted can expect to see their fathers over 8 days more each year than similar children in states that had not yet adopted such a policy.

Taken together, these results identify very specific policies that alter father–child contact. However, such models do not provide insight into whether future policies aimed at the establishment and collection of child support would produce similar effects. It is possible that increasing links between the father and child through child support transfers strengthens their relationship and fosters visitation. However, it may be other incidental effects of these policies that lead to contact. To further examine this issue we estimate structural IV Tobit models of the effect of establishing child support awards and receiving child support payments on father–child contact.

6.2. *Structural model*

The estimated coefficients from the GMM Tobit model with potentially endogenous explanatory variables, corrected for heteroscedasticity are presented in Table 4.[13] Again we find little evidence that contact is determined by characteristics of the mother and child. As in the reduced form models, lower levels of

Table 4. Structural tobit regression of annual father–child contact for children born to unmarried parents.

Variable	Father-child contact Tobit coefficients
Age of child (years)	−5.58* (2.98)
Male child	−4.85 (5.99)
Hispanic (mother)	−11.29 (10.28)
Black, non-hispanic (mother)	−3.77 (74.29)
Mother's education (highest grade completed)	1.22 (1.55)
Number of nonmarital children in household	−2.35 (3.24)
Mother currently married	−8.15 (10.26)
Mother teen at time of birth	−4.72 (7.17)
Father lives in the same state as child	59.77*** (10.58)
Father lives in the same city/county as child	28.63*** (10.05)
Mother born outside the USA	−3.04 (13.19)
Any child support agreement in effect	27.07*** (9.74)
Received some child support in past year	71.37*** (9.78)
Birth year trend	−4.41 (2.80)
Sample size	2189

Notes: Reported Tobit coefficients are for the latent variable. Marginal effects reported in the text are for the observed number of days of annual contact. Standard errors of the Tobit coefficients adjusted for clustering at the household level are reported in parentheses. Regression also includes a constant and indicators for missing mother's place of birth and missing father's residence.

*: $p < .10$, **: $p < .05$, ***: $p < .01$.

contact are evident for families with more non-marital children and for children of Hispanic mothers, but these coefficients are smaller and less precisely estimated in our structural model. In this model, contact significantly diminishes as the child ages. This effect may be due to estrangement between father and child as the time spent apart increases, or perhaps is caused by changes in the activities or preferences of the child. The positive relationship between proximity and contact remains.

In addition to the effects of family characteristics, our results indicate that child support plays a substantial role in the level of contact between fathers and their non-marital children. Though the presence of a child support award is associated with a small increase in father–child contact (nearly 10 more days per year), the effect of child support receipt is striking. Fathers who pay any child support will see their children 25 more days per year than children whose fathers pay no support. This effect is slightly larger than the increase in contact that would occur if an out-of-state father moved to the state in which his child resides.

The GMM estimation technique gives rise to a test statistic for the hypothesis that child support agreement and child support received are exogenous, i.e., not correlated with the disturbance. The use of overidentifying instrumental variables in addition to the potentially endogenous variables in this case leads to a test statistic distributed chi squared with 11 degrees of freedom. The resulting test statistic is 14.0, which is not statistically significant even at 20 percent. The result is that exogeneity is not rejected. This finding implies that any correlation between child support award or receipt and father–child contact is not caused by unobservables. Because we cannot reject the exogeneity of child support awards and receipt, our results can be interpreted to mean that policies that result in child support awards or payments by fathers who otherwise would not pay, can substantially increase contact between children and their fathers.[14]

Although not explicitly used in the GMM estimation reported above, the marginal probabilities from the child-support award and child-support receipt probit equations are reported in Table 5 to give the reader an idea of the factors that predict child support (and, especially, the effect of the policy instruments). Blacks and Hispanics are less likely either to have been awarded or to receive child support. As mother's education increases, the probability of both award and receipt rise. Older children are more likely to have an award and to receive child support, and there is a positive trend in child support awards. The likelihood of a child support award and receipt are higher when the father and child live in the same state, but being in the same city is unimportant. If proximity is correlated both with the desire to pay child support and the desire for contact, there may be a potential endogeneity problem by including father's residence in this model. However, our results imply that this may not be a problem, because living in the same city is correlated with father–child contact, but it is not correlated with child support payments. The fact that living in the same state is correlated with child support can be explained by the difficulty of enforcing child support laws on fathers living in another state (Peters et al., 1993).

Table 5. Probit regressions of child support agreement and child support receipt for children born to unmarried parents.

	Marginal Probability	
Variable	Any child support agreement	Any child support receipt
Age of child (years)	0.091*** (11.37)	0.027*** (2.67)
Male child	−0.010 (0.63)	−0.006 (0.36)
Hispanic (mother)	−0.129*** (4.75)	−0.116*** (3.89)
Black, non-hispanic (mother)	−0.076*** (4.08)	−0.079*** (3.96)
Mother's education (highest grade completed)	0.012*** (2.72)	0.012*** (2.64)
Number of nonmarital children in household	−0.063*** (2.72)	−0.061*** (8.03)
Mother currently married	0.096*** (3.62)	0.040 (1.47)
Mother teen at time of birth	0.008 (0.43)	−0.018 (0.92)
Father lives in the same state as child	0.248*** (10.62)	0.241*** (9.48)
Father lives in the same city/county as child	−0.011 (0.49)	−0.001 (0.06)
Mother born outside the USA	−0.055 (1.57)	0.011 (0.28)
Birth year trend	0.062*** (7.80)	0.010 (0.92)
Genetic testing laws	0.028 (0.86)	−0.036 (1.08)
Paternity ratio	0.011 (0.25)	0.011 (0.22)
Child support guidelines	0.071** (2.12)	0.087*** (2.59)
AFDC benefits[a]	−0.001 (0.24)	0.012** (2.04)
CSE administrative expenditures[a]	0.032** (2.22)	0.062 (0.71)
CSE collections to expenditures[a]	0.023*** (4.41)	0.016*** (2.68)
UIFSA[a]	0.016 (0.50)	0.066*** (2.59)
Sample Size	2189	2189
Chi-Squared Statistic (7 degrees of freedom)	29.780***	25.919***

Notes: Regressions also include a constant and an indicator for missing mother's place of birth. T-statistics calculated from standard errors adjusted for clustering at the household level are reported in parentheses.

[a]Measured at the time of birth for agreement equation and at time of interview for receipt equation; *: $p < .10$; **: $p < .05$; ***: $P < .01$.

State child support and welfare policies do affect the probabilities of having a child support award and of receiving child support. Specifically, awards and receipt are more likely in states that devote more resources to their CSE efforts and in states with high ratios of CSE collections to expenditures. Similar to findings in other work (Argys, Peters, and Waldman, 2001) the presence of child support guidelines also increases the probability that child support is awarded and received. Calculations using the results from Tables 4 and 5 suggest that when states adopted child support guidelines and UIFSA, the probability of child support receipt increased by 15.3 and 7.1 percent more children had child support awards. This, in turn would increase father–child contact by 4.5 days per year, a 16 percent increase in contact on average. Similarly, the doubling of each state's child support collection ratio will increase the probability of receiving child support by 6.2 percent and the incidence of child support awards by 8.9 percent. These increases in child support would translate into an increase in contact for all children with non-resident fathers of 1.4 days per year.

Though only selected instruments are individually significant, the sets of instruments are jointly significant predictors in both equations. Joint significance of the instruments makes the selected policy variables good candidates as instruments, but as discussed above, the instruments must also be uncorrelated with the disturbance in the contact equation. By designating CSE expenditures and CSE collections to expenditures as justifiably excludable instruments, we can test the excludability of all the other instruments. As discussed earlier, one could make a plausible argument that paternity policies such as genetic testing may independently affect father–child contact. However, our tests indicate that the exclusion restrictions cannot be rejected.[15] That is, that the paternity policies do not have an independent effect on father–child contact. The results of this structural model, combined with the results from the reduced-form models, suggest that policies aimed at improving a non-resident father's financial ties to his children have significant effects on father–child contact.

7. Conclusion

This paper uses data from the Survey of Income and Program Participation on children born to unmarried parents between 1985 and 1997 to investigate the relationship between policies aimed at providing financial support for children and contact between children and their non- resident fathers. Our reduced-form models, which measure the total effect of specific policies on contact indicate that, although not aimed at increasing father–child contact, more generous welfare benefits, efficient child support collection by the state Child Support Enforcement Office and interstate enforcement efforts do lead to greater visitation between children and their non-resident fathers.

Will new policy efforts aimed at increasing child-support awards and payments further alter the amount of contact that non-residential fathers have with their non-marital children? Reduced-form models don't provide the answer. To improve our understanding of the path through which these policies affect father–child contact, we examine the relationship between child support awards and payments and contact. Since it is likely that both of these behaviors are, in part, determined by unobservable characteristics of the father, we estimate an instrumental variables Tobit model. The model is identified by our assumption that child support policy variables can impact child support awards and payments, but father–child contact is not directly legislated.

Our results suggest that there are unintended, but desirable effects of child support establishment and collection activities that operate through their effects on child support awards and payments. Efficient collection policies and efforts to establish awards and adjudicate cases with out-of-state parents will not only increase the financial support available to children of unmarried parents, but, through their impact on child support agreements and payments, these policies will increase visitation and contact between these children and their fathers. Specifically, the presence of a child support agreement increases father–child contact by 10 days per year while receipt of child support leads to a substantial increase in contact of more than 25 days per year.

While these results suggest some positive unintended effects of child support and welfare policies on father–child contact, we have no evidence of the effect of such policies on other aspects of a non-resident parent's relationship with his child. As data sets which collect more detailed information on the quality of parent–child interactions become available, this line of research may be extended to examine the effect of such policies on other outcomes.

Notes

1. See http://www.sipp.census.gov/sipp/ for more information about SIPP.
2. We begin with a sample of 6101 children born between 1985 and 1997, who are living with their mothers, and who have an absent living father. Of that number, 2623 are identified as having parents who were divorced or separated, and 3430 are classified as being born outside of marriage. An additional 48 are not classified, because of missing or inconsistent data about marital dates. We also eliminate 1199 observations that are missing information about either child support receipt or father-child contact and an additional 42 observations with missing data on the socio-demographic variables. Children eliminated because of missing data are somewhat more likely to be older, African–American or Hispanic, come from larger families, have less educated mothers, and have fathers who live in different cities and states. We control for these observable characteristics in our analyses, but we do not know if the excluded children are different in unobservable ways.
3. For multiple-child families with awards, the respondent is first asked whether all children covered by the award see their father the same amount. If the answer is yes, contact information is available for all children. If contact is not uniform across children, this information is gathered only for the oldest child. If there is no award and all children have the same father, contact is only collected for the youngest child, but if children have different fathers, the data are collected for both the youngest and oldest child. In addition, contact information is not collected for children who have an award that is not the most recent award (e.g., from a different father) or if data about the type of award is missing. Of the 3430 children born outside of marriage, 7 percent are missing contact because of missing data on type of award or the award not being the most recent; 2 percent are missing because they have an agreement, but the children see the father different amounts of time and the data are collected for their sibling; 8 percent are missing because the respondents did not provide the information; and 12 percent are missing because the child had no agreement, and the data were collected for their sibling.
4. The percentage of *children* with missing contact and child support data is 34 percent, while the number of *families* with missing child support or contact information is only 20 percent.
5. Another significant source of missing data is incomplete migration histories. We use the child's state of birth to link the data on individual children with relevant state policies. We derive that information from the migration history. Because the migration history is sometimes incomplete, we are more likely to having missing information about state of residence for children living in families with frequent moves and for children who are older. To reduce the number of missing cases, we use the first state observed within three years of the child's birth. We used information on state of residence after the child's birth for 18 percent of the children in our sample. For 16 percent, of our sample, there was no reported state of residence within three years of the child's birth. For these children, we assigned the national population-weighted average of the policy variables.

6 To increase the prevalence and amount of child support received from absent parents, federal and state governments adopted a number of other child-support policies. To improve child-support collection, states subjected child-support orders to immediate income withholding and, in non-paying cases, states and local child support agencies imposed liens on bank accounts and other property, revoked drivers' and professional licenses and used a wide array of other tools to collect child support. In preliminary unpublished work, after controlling for CSE expenditures and collections, we found only a weak relationship between these tools and the likelihood of child support awards and receipt among unmarried women. Therefore, we do not include indicators of the adoption of these policies in our analysis.
7. These remaining 14 states all adopted UIFSA in 1998.
8. This statistic does not exactly measure the fraction of births in a year in which paternity is established, because some of the paternity cases may be for older children. However, it is the standard measure of state paternity establishment effectiveness (OCSE, various years).
9. We need to assume that two policy variables meet the criteria, because our model includes two potentially endogenous variables.
10. A basic reference on GMM is William Greene (2000, Section 11.5, pp. 474–491), and a somewhat more advanced reference on non-linear IV estimation is Davidson and MacKinnon (1993). The Tobit model is discussed in detail in Greene (2000, Chapter 20).
11. We convert the Tobit coefficients into the marginal probability on observed days of contact by multiplying the coefficient reported in Table 4 by the proportion of the sample with non-zero values for days of contact (.360)
12. Because both policies measured at the time of birth and current policies could potentially have an impact on current contact, we also estimated a specification that included all of these policies in the same model. However, problems of colinearity reduce the individual precision of the coefficients in that model, and we chose to report the two models with policies at different times entered in separately.
13. We also ran separate models by race. Because those results do not differ much from what is presented in Table 4, we do not report estimates from the race specific models.
14. Although our test statistics reject the hypothesis that child support and awards are endogenous, we also estimated the model jointly treating child support award and receipt as endogenous. These results (available from the authors upon request) indicate that receipt of child support does not have a significant impact on contact. The presence of a child support award is associated with a significant increase in annual father-child contact of 47 days. This estimate is less precisely estimated than the effect reported in Table 4.
15. The test statistic indicates that we cannot reject the exclusion of the instruments ($p = .23$). We re-estimate the model excluding the two instruments most closely correlated with the disturbance, AFDC payments at the time of birth and at the time of the child support interview. The structural Tobit results on amount of father–child contact are essentially unchanged by the elimination of these instruments.

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