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Testing the Economic Independence Hypothesis: The Effect of an Exogenous Increase in Child Support on Subsequent Marriage and Cohabitation

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Abstract We examine the effects of an increase in income on the cohabitation and marriage of single mothers. Using data from an experiment that resulted in randomly assigned differences in child support receipt for welfare-receiving single mothers, we find that exogenous income increases (as a result of receiving all child support that was paid) are associated with significantly lower cohabitation rates between mothers and men who are not the fathers of their child(ren). Overall, these results support the hypothesis that additional income increases disadvantaged women's economic independence by reducing the need to be in the least stable type of partnerships. Our results also show the potential importance of distinguishing between biological and social fathers.

Keywords Child support · Cohabitation · Independence hypothesis · Marriage · Single mothers

Introduction

The potential impact on marriage of policies that increase economic resources is a topic of longstanding interest for demographers and policy analysts (see, e.g., the classic paper by Hannan et al. 1977). However, the relationship between economic resources and partnering behavior may be changing (Sweeney 2002). Well-known trends that affect the lives of children—our focus here—include substantial increases in cohabitation (Kennedy and Bumpass 2008), declines in marriage (Cherlin 2009), new understanding of the connections between economic resources and relationship instability in

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many "fragile" families (Carlson and England 2011; McLanahan 2011; McLanahan and Beck 2010), and a new awareness of different types of family complexity (Carlson and England 2011; Cherlin 2009; Thomson and McLanahan 2012). One type of family complexity that may be particularly consequential for children is mothers who live with a "social father" (i.e., a man who has not fathered any of her children); some research suggests that social fathers are less involved with children than biological fathers (Berger and Langton 2011). In this article, we reexamine the relationship between economic resources and partnering, using more recent data and considering not only marriage but also cohabitation—and, in particular, whether cohabitation is with a biological father or social father.

These issues are particularly consequential for economically disadvantaged mothers and their children. An additional dollar of income is likely to have a greater effect for those with lower initial resources. Moreover, economically disadvantaged mothers are the focus of much social policy. Part of the impetus for some of the more radical changes in U.S. welfare policy has been a sense that the old welfare system promoted marital dissolution, increased nonmarital births, and enabled unmarried mothers to remain single (for a review of the mixed empirical results, see Moffitt 1998, 2003).[1] Because of the concern that programs that provide more economic resources inadvertently promote single mothering, much recent U.S. policy aims to encourage marriage among single mothers. In this context, the extent to which a mother's economic resources influence her marriage and cohabitation decisions has obvious implications for social policy.

As discussed later, any analysis of economic resources and partnering faces challenges in determining whether the relationship is causal and in identifying its direction. In this article, we take advantage of a policy experiment that resulted in randomly assigned differences in the amount of child support received among welfare participants to investigate the effects of economic resources (in this case, from one source: child support) on cohabitation and marriage. Because the differences in child support received when it was paid are unrelated to any other factors, we can estimate a causal effect of these additional economic resources. This research thus updates the prior research on the effects of income on partnering, focusing on a particular type of income (child support) among disadvantaged mothers and using superior methods and a nuanced measure of partnering outcomes (marriage and cohabitation with biological and social fathers).

Prior Literature

Although substantial literature in sociology and economics has examined the effect of income on union formation and family structure, determining a causal relationship is complicated (for a review of the empirical research and the estimation difficulties, see

[1] For example, the "Findings" section of the 1996 Personal Responsibility and Work Opportunity Reconciliation Act (PL104-193) notes that, "The increase in the number of children receiving public assistance is closely related to the increase in births to unmarried women. . . . The negative consequences of an out-of-wedlock birth on the mother, the child, the family, and society are well documented." It concludes that, "Therefore, in light of this demonstration of the crisis in our Nation, it is the sense of the Congress that prevention of out-of-wedlock pregnancy and reduction in out-of-wedlock birth are very important Government interests and the policy contained in part A of Title IV of the Social Security Act (as amended by section 103(a) of this Act) is intended to address the crisis."

Burstein 2007). One critical difficulty is that individuals with different levels of income may be different in a variety of unmeasured ways that are correlated with family structure. For example, those who have difficulty maintaining commitments may be more likely to transition in and out of relationships, and they may also be more likely to transition in and out of employment, leading to lower overall income. A second issue is that income and family structure are likely to have reciprocal relationships: family structure clearly affects the resources available to an individual, and resources can affect family structure. Third, partnering reflects a matching process involving at least two actors; even if a given level of income increases an individual's desire to be in a partnership, there may not be an available and willing partner. The number of available "marriageable" partners may affect union formation (e.g., Raley 1996). Finally, many studies of the effects of income have examined the effects of earnings (rather than the effect of income *per se*), which complicates the analysis because earnings levels may reflect labor supply decisions associated with the expected stability of unions (Özcan and Breen 2012), and the working conditions of the jobs typically available to disadvantaged women may affect family formation (Joshi et al. 2009).

In this article, we are particularly interested in the effect of income on women who are already single mothers, considering whether they marry, cohabit, or stay single—and, if they cohabit, whether it is with a man who is the biological father of at least one of her children. In the sections that follow, we summarize three categories of recent research that focuses on (1) the effects of income on transitions from being unpartnered to marriage and cohabitation, (2) the effects of income on transitions from cohabitation to marriage, and (3) the effects of income on the type of cohabitation (with a biological or social father). Within each section, we review available theory; and because we want to pay particular attention to research regarding low-income mothers, we discuss research on welfare and child support income in particular, as well as income in general.

Effect of Income on Transitions to Partnerships

Theory is ambiguous regarding the likely effects of economic resources on entering a romantic partnership, whether cohabitation or marriage. The most common approach assumes an independence effect: because additional income reduces the need for a woman to find a partner with whom she can pool resources, income is assumed to make her more independent and less likely to partner (Burstein 2007; Bzostek et al. 2012; Carlson et al. 2004b; Cherlin 1992).[2] Alternatively, increases in income may lead to increases in cohabitation or marriage if they make a woman a more attractive partner (e.g., Carlson et al. 2004a; Oppenheimer 1997). Finally, there may be no relationship: small amounts of additional income may not matter if a woman is receiving income-tested welfare benefits and thus loses benefits with additional income. Moreover, even given a relationship between income and marriage, there may be no relationship for

[2] A closely related, but distinct, argument suggests that women's increasing opportunities in the paid labor market lead to lower rates of marriage because the economic gains from specializing (i.e., the potential gains from allowing, for example, the man to work more hours in the market and the woman to concentrate on home production) have been eroded (Becker 1981). The economic independence argument focused on specialization is less relevant to our interests in the effects of an increase in child support or other unearned income—a change that is not directly related to wages.

cohabitation; some researchers argue that the factors related to the transition to cohabitation are substantially different from the factors related to the transition to marriage (e.g., Reed 2006).

Several authors have concluded that although the economic independence arguments have some theoretical support, the empirical evidence is weak (e.g., Burstein 2007; McLaughlin and Lichter 1997; Oppenheimer 1997; White and Rogers 2000). The argument suggests that those with the fewest resources should be most likely to be partnered; this is not the pattern typically observed in straightforward comparisons (e.g., White and Rogers 2000) or in more sophisticated empirical analyses (e.g., Aassve 2003; Carlson et al. 2004a; Clarkberg 1999; Gibson-Davis 2009; Goldscheider and Sassler 2006; Joshi et al. 2009; Lichter et al. 1992; Raley 1996; Sassler and Schoen 1999; Schneider 2011; Sweeney 2002; Thornton et al. 1995; Xie et al. 2003). The research on entries into unions in general (often treating marriage and cohabitation as competing risks) and research focusing specifically on entering cohabitation generally shows a weak relationship between economic resources on these transitions; when there are associations, men's resources seem more important than women's (e.g., Carlson et al. 2004a; Clarkberg 1999; Goldscheider and Sassler 2006; Raley 1996). Moreover, with respect to marriage, a recent review (Burstein 2007) concluded that studies of individuals find either that women's resources are not associated with the transition to marriage or that women with higher earnings, education, or income have an increased likelihood of marriage, in contrast with the expectations of the independence hypothesis. However, few of these studies have attempted to control for unmeasured factors.

A substantial literature focusing on lower-income mothers has examined the effect of a particular income source—specifically, welfare income—on marriage. Welfare income may have similar effects to other types of income, an independence effect that makes it less likely to partner, and a marriage market effect that makes it more likely to partner. However, the effect of welfare income is complicated because those who marry are less likely to be eligible for some programs. A typical approach uses variations in state benefit levels to see whether transitions to marriage were more likely or less likely in states with more generous benefits (Moffitt 1998; for a review, see Moffitt 2003). This research has generally found no effects or very small effects of welfare policy (although for an alternate finding, see Knab et al. 2009). In another line of research, evaluations of several welfare reform initiatives have tested whether changes to the amount of welfare benefits and other changes to welfare rules have affected marriage rates for single mothers using random-assignment designs, which should control for the unmeasured variables that have been a limitation of most previous work. A meta-analysis of 14 of the random-assignment studies (Gennetian and Knox 2003) found no overall effects of these reforms on marriage, nor did they find consistent effects on marriage among subgroups. Part of the reason for limited effects may be that the reforms typically combine many changes, some of which may have increased marriage rates and some of which may have decreased marriage rates. Although the welfare reform studies have generally shown no effects on marriage, some individual studies—especially those that provided more income—have found effects, either overall or among some subgroups (see Gassman-Pines and Yoshikawa 2006; Gennetian et al. 2005; Harknett and Gennetian 2003). The effects, with those receiving more income being more likely to marry, were particularly large among women who were most disadvantaged (Gennetian and Knox 2003). Putting together the experimental research with a variety of nonexperimental studies leads to an overall conclusion that the effects of welfare programs on marriage or cohabitation are likely small (Burstein 2007).

A much smaller literature has explored the effects of child support income on transitions to marriage and cohabitation. Carlson et al. (2004a) argued that child support enforcement has potentially competing effects: child support, like any other income source, makes low-income single mothers more able to live independently, but it also could make them more attractive in the marriage (or cohabitation) market. An additional complication is that having to pay child support increases the incentives for fathers to live with the mothers of their children, so strong enforcement of child support may be associated with increased marriage to a father. Carlson and colleagues (2004a) found in some empirical models that couples are more likely to be married if child support enforcement is higher, and this result is consistent with other studies finding that more child support enforcement is associated with lower divorce rates (Nixon 1997) and increased marriage among those who are already parents (Acs and Nelson 2004). However, Carlson and colleagues' (2004a) findings were not consistent across models. Similarly, Knab et al. (2009) found no effect of child support policies on the likelihood of marriage among couples who already had a nonmarital birth together but who were not cohabiting at the time of the birth. One review concludes that the findings on the effects of child support on marriage and divorce are quite mixed (Pirog and Ziol-Guest 2006), reflecting the absence of consistent effects across models found by some authors (Carlson et al. 2004a; Knab et al. 2009) as well as findings that child support is associated with increased marriage (Acs and Nelson 2004; Nixon 1997), no effect on marriage (Heim 2003), and decreased marriage (Knab et al. 2009).

Effect of Income on Transitions From Cohabitation to Marriage

Some studies have examined the relationship between income and the transition to marriage among women currently cohabiting (e.g., Brown 2000; Duvander 1999; Gibson-Davis 2009; Manning and Smock 1995; Sassler and McNally 2003; Smock and Manning 1997; Wu and Pollard 2000). Economic resources could affect this transition because married couples are treated differently from cohabiting couples by the tax and transfer system; however, whether a couple would have more or fewer resources were they to marry differs widely across couples, making simple statements about the general effect of marriage difficult (Burstein 2007). Recent qualitative research has emphasized that higher income may provide the economic resources that some couples believe are necessary to marry, rather than cohabit (Edin 2000; Gibson-Davis et al. 2005; Smock et al. 2005), an issue that is likely to be particularly salient for lower-income women.

In an early study, subsequently confirmed by a number of additional quantitative studies, Smock and Manning (1997) showed no effect of the economic resources of a cohabiting woman on her transition to marriage; however, cohabiting men with higher earnings are more likely to marry. Similarly, a woman's employment status is not related to the transition to marriage among low-income cohabiting women (Lichter et al. 2006).

Turning to research on particular sources of income, welfare benefits have no relationship to the rate of marriage among cohabitors with low income (Lichter et al. 2006). The experimental studies, even those that have found some effects of welfare on marriage, do not show that increased welfare income for women already cohabiting is associated with transitions to marriage (Gennetian 2003). Because child support is seldom pursued when two parents cohabit or are married, the level of child support enforcement should not have a large effect on the transition to marriage among couples who are cohabiting. However, Knab et al. (2009) found that more child support enforcement was associated with a lower likelihood of transitioning from cohabitation to marriage. Looking at the research as a whole, we find little quantitative evidence to date that a cohabiting woman's resources, either in general or of a particular type, have a strong effect on the likelihood of marriage.

Effects of Income on Type of Cohabitation

Cherlin and Fomby (2004) made an important contribution in distinguishing types of cohabitors, differentiating women cohabiting with a father of at least one of their children from those who are cohabiting with someone who is not a father of one of their children. In examining living arrangements for low-income mothers, they found in their baseline survey that cohabiting with a biological father is more common than cohabiting with a social father. By the second wave of their survey, though, cohabiting with a social father was more common. Cohabiting relationships of all types are unstable, regardless of whether they have produced children (e.g., Kennedy and Bumpass 2008). Relatively few researchers have followed the lead of Cherlin and Fomby in separating types of cohabitation between those with biological and those with social fathers (although see, e.g., Manning and Brown 2006), and we are not aware of recent research that examines the relationship between income and these types of cohabitation.

Summary

Reviewing the theoretical and empirical research, the effect of economic resources on partnering has clearly been an important topic for researchers across multiple disciplines. However, research on the effect of income on the partnering of disadvantaged single mothers, while a central focus of policy debates, has received less focused attention. Much of the extant literature does not fully differentiate between lower-income and higher-income women, or does not differentiate between those who are already mothers and those who are not. Moreover, much research is focused on the effects of income on union dissolution rather than formation; much of the policy discussion, in contrast, focuses on those who are already single parents. Perhaps most importantly, the current literature is limited in its ability to disentangle factors that may be related to both partnering and economic resources. Finally, although some literature has explicitly considered cohabitation, little research has followed Cherlin and Fomby (2004) in trying to differentiate types of cohabitation. This article reports the results of an analysis of outcomes for low-income mothers exposed to a unique random-assignment experiment that allows us to address some of these limitations. Before presenting our analysis, we discuss the context of this experiment.

Policy Context

Because single-parent families have historically been economically vulnerable, several government programs have been designed to increase their incomes. Most public attention has focused on the traditional cash welfare program, Aid to Families with Dependent Children (AFDC), which was replaced by Temporary Assistance for Needy Families (TANF) in the welfare reform of 1996. Welfare reform was influenced by a sense that the old income-support scheme, in which eligibility for many supports was restricted to single-parent families, encouraged separations and nonmarital births because more resources were available to single-parent families than to two-parent families. Similarly, concerns existed that programs such as AFDC, which gave additional resources to those with larger families, encouraged additional births, particularly nonmarital births. TANF responded to these issues by (1) allowing states to choose to provide the same benefits regardless of the number of adults in the home, (2) giving states the freedom to keep benefits constant regardless of the number of children, and (3) providing incentive payments to states that reduced nonmarital births. Other provisions of TANF also responded to criticisms of AFDC: because federal programs were seen as too inflexible and bureaucratic, states were given great freedom to design their own TANF programs. In addition, because providing income for an unlimited period of time was seen as serving to keep single mothers out of the labor force, TANF programs were time-limited and focused on requiring work in exchange for benefits.

The other major social policy affecting single-parent families is child support. When children live with only one biological parent, the nonresident parent (usually the father)[3] is typically required to pay child support to the resident parent (usually the mother) to contribute to child-rearing expenses. A public child support agency in every state helps to locate nonresident parents and establish child support orders, collects child support paid primarily through wage withholding, and then distributes that support to the resident-parent family. Child support orders typically do not change if a single parent (re)marries or begins to cohabit with a new partner.[4]

Child support and TANF are strongly related; in fact, early child support policy was designed primarily as a tool to recoup (or prevent) welfare expenditures on children (Garfinkel et al. 1998). The 1996 welfare reform that created TANF also allowed states to set their own rules for what happens when child support is paid for a family receiving TANF. In most states, any child support paid by the nonresident father when the mother is receiving TANF is retained by the government and used to defray the costs of public assistance, rather than being distributed to the family (Cancian et al. 2008). Some states instead maintained the pre-reform policy of allowing the resident mother to receive the first \$50 of child support paid each month (a \$50 "pass-through") and retaining only amounts above that limit (Cancian et al. 2007). One state—Wisconsin—has adopted a unique policy, allowing most mothers to keep all the child support paid by the father but without that support affecting the level of cash welfare received by the mother and her family. This policy was implemented as part of the state's TANF program,

[3] For simplicity, in this article, we use gendered language and refer to nonresident parents as "fathers" and resident parents as "mothers." Child support policies and income support policies are not explicitly gendered, and rights and responsibilities are assigned by residential status. However, particularly among low-income families (including those participating in welfare programs), most resident parents are mothers, and most nonresident parents are fathers (Grall 2011).

[4] For recent research on changes to child support orders, see Ha et al. (2010). In the United States, child support orders are typically based on the personal income of the biological parent, not their family income (Gold-Bikin and Hammond 1994); and, in contrast to some other countries, child support orders generally do not differ based on the family status of either parent (Meyer et al. 2012; Skinner et al. 2007).

Wisconsin Works (W-2). The child support component of W-2 was the subject of an experimental evaluation, with a random sample of TANF recipients and applicants assigned to a control group and receiving only partial child support (the first $50 per month or 41 % of the amount paid, whichever was more) whenever they received TANF, and those in an experimental group receiving all current child support paid, regardless of TANF receipt. The evaluation was in effect until July 2002, when all cases began to receive all child support paid on their behalf.

We use the variation in child support received by the control and experimental groups that was caused by random assignment to identify the effect of child support income on subsequent marriage and cohabitation. This strategy depends on the successful implementation of random assignment, which was documented as part of the original Child Support Demonstration Evaluation technical report (for details, see Cancian et al. 2001). Families were assigned to the experimental and control groups based on a randomly distributed digit of the mother's Social Security number. Families in the experimental group received more child support as a direct mechanical effect of the policy change, but assignment status was consequential only for those receiving TANF cash benefits and child support in the same month. TANF participation tended to be short, and child support receipt was irregular. Nonetheless, although the difference in income is fairly modest, it provides an opportunity to test the causal effects of an exogenous increase in family income.

A random-assignment experiment provides an estimate of the total effect of an increase in child support income, but the effect could be realized through a number of pathways. For example, if a mother received increased child support income, it could make her less desperate to find someone with whom she can share expenses, or it could increase her sense of control over her life—and either outcome could make her more selective in partnering. The experiment offers a distinct advantage for estimating whether income has an effect but does not allow us to identify the particular pathway (e.g., whether the effect comes through less economic need or increased sense of control).

In summary, the random-assignment experiment means that the two groups of low-income single mothers should be identical except that one group receives more child support income than the other. This enables us to contribute to the debate about whether income causes changes in partnering behavior (and thus living arrangements) for this vulnerable population.

Data, Sample, and Methods

Our sample includes mothers who entered the Wisconsin TANF program (W-2) in its first 10 months of operation, between September 1997 and July 1998. We use merged administrative records drawn from the welfare, child support, and Unemployment Insurance systems, merged through Social Security numbers (Cook and Brown 2001). We also use information available for a random sample of the control and experimental groups from the Survey of Wisconsin Works Families (SWWF).

Our primary sample includes mothers who responded to the third wave of the SWWF (n = 709), which included measures of family structure and living arrangements. The original survey sample was drawn from the W-2 administrative records and is linked to it (Krecker 2001). Of the original sample drawn for the survey, 91 %

responded to at least one of the two first waves of the survey and were thus eligible for the third wave; for the third wave, a random sample of one-third of those eligible was selected, and 82 % responded (Krecker 2005). The third wave was collected in the spring and summer of 2004. Using detailed administrative data available for both respondents and nonrespondents, weights were created to account for nonresponse (see Ziliak 2004) and are used here. Because our primary interest is those who were not married when they entered W-2 in 1997–1998, we exclude 35 mothers who the administrative records show were married and two mothers with missing marital status at that time, leaving a final sample of 672. We do this because we would like to differentiate the effect of income on *getting* married from that of *staying* married, and there are too few cases who were married when they entered W-2 to analyze separately.[5] Because the administrative data on cohabitation are unreliable, our base sample includes both those who were single and those living with a cohabitor.[6]

Table 1 reports basic demographic characteristics for mothers at the point when they entered W-2 as well as their recent work, welfare, and child support history. The first set of columns shows the results for the full sample, followed by results for the experimental and control groups separately. As expected, given successful random assignment, there are no significant differences in the distribution of baseline characteristics of the experimental and control groups, nor were there differences in prior welfare use, prior employment, or prior child support. Most mothers were in their 20s. Almost two-thirds of mothers in the full sample were black, and one-quarter were white. Many mothers had low levels of formal education: 54 % of the full sample had not completed high school, and only 10 % had more than a high school education. At the time they entered W-2, most mothers had one or two children, but 36 % of the full sample had three or more. Most women had at least one young child; only one-fourth of the full sample had no preschool children. In about one-fifth of all cases, the child support system official records of paternity show that the mother had had children with more than one father. Most mothers lived in Milwaukee County, the primary urban county in the state.

At entry to W-2, more than one-half the mothers were longer-term AFDC participants. In the two years before entering W-2, more than 80 % of mothers had some formal employment experience, as measured by earnings recorded in the Unemployment Insurance system; and 13 % of the full sample had worked in every quarter. Most women had a child support order in place at entry, although only 23 % had child support paid on behalf of their family in the prior year, and only 10 % of the full sample had at least $1,000 paid on their behalf in the prior year. Nonetheless, when child support was paid, it averaged about $100 per month. Overall, the descriptive statistics show a group of mothers who are disadvantaged on several dimensions.

[5] Including the 35 mothers who were married at entry would provide similar results. Although none of our main results that were even marginally significant change in sign, some results that were not marginally statistically significant become so with the larger sample. The most important is that in one of the subgroups—namely, those who did not receive AFDC in the prior two years—women in the experimental group were marginally less likely to be married ($p = .07$).

[6] If we were to exclude the 65 mothers who the administrative records show as not married but who were living with an unrelated male who was within 10 years of her age (a potential cohabitor), the main conclusions would be identical.

Table 1 Mothers' characteristics and program participation at entry to W-2

	Total		Experimental		Control	
	N	Weighted %	*N*	Weighted %	*N*	Weighted %
Age (p = .1611)						
<20 years	91	12.6	53	15.3	38	10.1
20–29 years	370	55.7	161	52.5	209	58.4
30–39 years	173	26.6	88	28.0	85	25.6
40 years and older	38	5.1	15	4.2	23	5.9
Race (p = .2000)						
White	196	24.1	84	21.2	112	27.0
Black	425	66.3	205	68.1	220	64.6
Hispanic	30	5.8	15	5.6	15	6.0
Other	21	3.8	13	5.2	8	2.4
Education (p = .7490)						
Less than high school	338	53.6	165	55.0	173	51.9
High school	257	36.8	119	35.8	138	38.1
More than high school	77	9.6	33	9.2	44	10.0
Number of Children (p = .1129)						
0 (pregnant at entry to W-2)	11	0.9	4	0.6	7	1.3
1	249	34.9	120	35.5	129	34.2
2	187	27.9	76	24.2	111	31.4
3 or more	225	36.3	117	39.7	108	33.1
Age of Youngest Child (p = .3211)						
0–2 years (includes those pregnant at entry)	401	57.8	190	58.9	211	56.9
3–5 years	115	17.7	59	19.1	56	16.5
6–12 years	127	20.3	52	17.2	75	23.1
13–17 years	29	4.2	16	4.8	13	3.6
Number of (legal) Fathers (p = .7899)						
0	182	25.0	86	19.7	96	24.3
1	369	56.0	169	54.6	200	57.4
2	121	19.0	62	25.7	59	18.3
County (p = .5345)						
Milwaukee	474	76.3	226	77.0	248	75.8
Other Urban	121	15.3	51	13.8	70	16.6
Rural	77	8.4	40	9.2	37	7.7
Number of Months of AFDC Receipt in Prior 24 months (p = .5017)						
No AFDC	150	13.5	66	12.7	84	14.3
1–18 months	238	32.4	110	30.8	128	33.8
19–24 months	284	54.1	141	56.4	143	51.9
Quarters Employed of the 8 Quarters Before Entry (p = .1230)						
0 quarters	84	15.1	38	14.7	46	15.5
1–4 quarters	250	43.1	107	38.7	143	47.2
5–7 quarters	215	29.4	112	33.5	103	25.5

Table 1 (continued)

	Total		Experimental		Control	
	N	Weighted %	N	Weighted %	N	Weighted %
8 quarters	123	12.5	60	13.1	63	11.8
Child Support Order/Payments on Behalf of Family in Prior 12 months (p = .7061)						
No order	281	41.3	131	42.3	150	40.5
Order but no child support paid	237	36.1	111	35	126	36.9
$1–999 CS paid	82	12.5	42	13.6	40	11.4
$1,000 or more CS paid	72	10.1	33	9.1	39	11.2

Notes: The p values for a Pearson χ^2 test of the difference between the experimental and control group are in parentheses.

Empirical Approach and Method

Our basic approach is to examine whether unmarried mothers randomly assigned to the experimental group (who received the full amount of child support paid on their behalf) have different partnering from those assigned to the control group (who received only a portion when they were receiving W-2). Any differences would provide evidence of a causal effect of income (in this case, child support) on mothers' partnering. To validate this analytic strategy, we must confirm that random assignment worked—that is, that individuals in the experimental and control groups did not differ from each other at the entry into the program except by chance. In addition to the comparison of descriptive statistics for our analytic sample in Table 1, a variety of analyses demonstrate that random assignment worked and that the experimental and control groups are equivalent.[7] Second, we must confirm that the intervention worked—that is, that the experimental group actually received more child support than the control group. As reported by Cancian et al. (2008), mothers in the experimental group received over 20 % more support in the first year, decreasing to 12 % more in the third year. This effect is due both to more support being paid on behalf of families in the experimental group (fathers being more likely to pay, and paying higher average amounts) and to mothers in the experimental group receiving more of the child support that was paid. Although the experimental effect declines over time as more control-group mothers move off welfare and begin to receive the full amount of child support paid on their behalf, the effects for those new to the welfare system are larger and grow over time.

[7] Cases receiving AFDC on August 31, 1997, were randomly assigned to the experimental or control group on that date; cases applying for W-2 were given a random-assignment code on application. Cases in the experimental and control groups entered the W-2 program at equivalent rates; that is, there was no differential diversion. Further, there was no difference in rates of entry to W-2 in the group in which differential diversion would have been most likely: among those with child support amounts of more than $1,000 in the previous year, experimental-group and control-group members entered at the same rate. An analysis of whether a wide variety of characteristics can be used to predict experimental-group status shows that only three of 33 variables are statistically significantly related ($p < .05$) to being in the experimental group, which is not many more than the number that would be predicted by chance. As described later, we control for the variables shown to predict experimental-group status. See Cancian et al. (2001) for more details.

Having established that those in the full child-support (experimental) group are not substantially different at entry from the control group but that they do receive more child support, we now turn to an examination of whether being in the full child-support group is related to partnering—that is, to marriage and cohabitation patterns. We examine living arrangements at the time of the third-wave interview, in 2004, about 6.5 years after random assignment. We divide mothers into four categories based on mothers' self-reports. At the time of the third-wave interview, 11 % of mothers were married, 72 % were single (no partner), 7 % were cohabiting with a father of one of their children, and 10 % were cohabiting with someone who is not a father of one of their children (a "social father").

For mothers who were cohabiting, we distinguish between fathers and nonfathers based on the report of the mother, regardless of when a child was born or whether paternity has been formally established. For those mothers who were married, we do not distinguish between being married to a father and nonfather for two reasons. First, because relatively few mothers were married, we do not have the sample size to support estimates that distinguish between husbands who were fathers and those who were not. Second, although biological and nonbiological fathers may behave differently (Hofferth 2006; Hofferth and Anderson 2003), marriage is associated with more stable unions (Bumpass and Lu 2000) and greater pooling of resources than cohabitation (Hamplova and Le Broudais 2009; Winkler 1997), and we expect biological ties to be less important to the distribution of resources within marriage than within cohabiting unions (Berger et al. 2008).

Given valid random assignment, the cases in the experimental and control groups are equivalent in all ways except for being exposed to a different policy regime. As a result, a simple comparison of partnership status could be used to test the effect of having more income. We nonetheless control for selected covariates in a multinomial logit regression because using control variables provides more precise estimates of the relationship between experimental group status and our outcome of interest. Control variables also account for any chance differences between the experimental and control groups. Control variables include mother's age, race, education, the age of her youngest child, whether the mother had more than $1,000 in child support paid on her behalf in the year prior to entering W-2, urban/rural status, and variations in the period of random assignment (which differentiates AFDC recipients at the time of transition from those who applied to TANF later).[8]

For ease of interpretation, we also use the multinomial logit regression coefficients to calculate regression-adjusted proportions of the various categories of partnership status among the experimental and control groups. We do this by calculating the predicted probability that an experimental-group mother will be in each category, using the coefficients from the model and the weighted mean for each control variable; then we repeat the calculation for the control group. We then compare the proportions of experimental- and control-group members in each living arrangement category, using a *t* test.[9]

[8] This list of control variables includes all those used to determine regression-adjusted differences between the experimental and control groups in the final evaluation of the child support reform (see Meyer and Cancian 2001: 29–30).

[9] If the means of control variables were equal in the experimental and control groups, the significance of the experimental group coefficient should be equivalent to the significance of the difference in the estimated proportions. Because there are small differences in the means of the control variables, the significance of proportions can be slightly different.

We are particularly interested in any differential effects of additional income for three subgroups. First, we estimate the effect of the policy on a subgroup for which child support income might be especially salient, given their own limited earnings potential: women with less than a high school education when they entered W-2. This subgroup may be particularly relevant in testing the hypothesized effect of economic independence: they might be most likely to partner for economic reasons, and thus additional resources might cause them to be more likely to be unpartnered.

A second subgroup of interest is mothers who had no history of AFDC receipt in the two years prior to entering W-2. We are particularly interested in this subgroup for two reasons. First, long-term welfare participants may have characteristics that make them unlike other single mothers, and thus examining new participants provides more generalizable results. Moreover, time limits on receiving benefits were instituted under welfare reform, which means that there will be very few long-term recipients in the future. Second, the experimental effect on child support receipts is larger for new participants, perhaps because they had no recent experience with having child support retained (Cancian et al. 2008). This may increase our ability to detect the effects of income on living arrangements among this group.

Our last subgroup of interest is mothers of younger children. We are interested in this group, first, because any changes in living arrangements will affect these children for a longer proportion of their childhood. Second, mothers caring for young children face high child care costs if they work, which may make them particularly responsive to changes in income that are unrelated to earnings. Finally, mothers of young children may be more likely than other mothers to be open to having an additional child, which may change their partnering patterns.

Results

Table 2 reports the basic experimental results; we primarily focus on the first row, showing the effect of belonging to the experimental group. Six years after random assignment, mothers who are able to receive the full child support paid (the experimental group) are somewhat less likely to be cohabiting with a social father (a man who has not fathered any of her children) ($p < .06$). Translating the coefficients into predicted probabilities, 6.7 % of mothers in the experimental group are predicted to be cohabiting with a social father, compared with 11.2 % of those in the control group. As shown in the table, there is no statistically significant difference between the experimental and control groups in either the probability of being married or cohabiting with the father of one of her children. Calculating the predicted probability of being in each group reveals that the omitted category—remaining single—does show a difference between the experimental and control groups: 79.3 % of the experimental group, compared with 72.9 % of the control group, is predicted to be single and not cohabiting ($p < .09$). Thus, the exogenous increase in income marginally reduced the proportion of mothers cohabiting with a man who is not the father of any of her children but increased the proportion who remained single.

The coefficients on the control variables are typically in the expected directions. African American mothers are significantly less likely to be married or to be in either of the cohabiting groups; thus, they are most likely to be unpartnered. Those with the

Table 2 Multinomial model of living arrangements, full sample

	Married		Cohabiting, Father		Cohabiting, Nonfather	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
Experimental Group	–0.247	0.287	–0.137	0.338	–0.586†	0.310
Child Support History in Year Prior to Entering W-2						
$1,000 or more paid on behalf of family	0.149	0.445	–0.666	0.548	0.070	0.420
Age >30 Years	–0.521	0.428	–0.691	0.495	–0.680†	0.376
African American	–1.183**	0.344	–1.106*	0.438	–0.891*	0.385
Education (vs. less than high school)						
High school	0.390	0.311	–1.060*	0.415	–0.361	0.350
More than high school	0.712	0.440	–0.728	0.688	0.811*	0.404
Region (vs. Milwaukee)						
Other urban county	–0.075	0.399	0.107	0.504	–0.282	0.458
Rural county	0.186	0.458	0.353	0.589	0.509	0.503
Age of Youngest (vs. <3 years)						
Age 3–5 years	0.718*	0.366	0.168	0.468	0.350	0.390
Age 6 or more years	0.278	0.456	–1.793*	0.792	0.489	0.382
Intercept	–1.340**	0.388	–0.877*	0.413	–1.213**	0.428
Log-Likelihood			–559.09			

Notes: Model also controls for period entering W-2, a proxy for long-term receipt; there are no statistically significant estimates. A Hausman-McFadden test suggests the independence of irrelevant alternatives (IIA) assumption has not been violated.

† $p < .10$; * $p < .05$; ** $p < .01$

oldest children are least likely to be cohabiting with a man who is a father, perhaps reflecting the instability of cohabiting relationships. Those whose youngest child was between ages 3 and 5 at entry into W-2 are more likely to be married. Those with less education and older mothers are less likely to be cohabiting with a social father. There are no detectable differences based on the level of child support before W-2 entry, nor are there differences between those living in Milwaukee County (the largest urban county in the state) and other regions.

To further test the hypothesized effect of economic independence, we estimate the effect of the policy on several subgroups. Given smaller sample sizes for the subgroups, some reduction in statistical significance might be expected. We begin by focusing on a group for which child support income might be especially salient, given their own limited earnings potential: women with less than a high school education when they entered W-2. As shown in Table 3, among those with less than a high school education, mothers in the experimental group are less likely to cohabit with nonfathers. Transforming the coefficients reveals estimates of 5.6 % of experimental-group mothers, compared with 11.2 % of control-group mothers ($p < .08$), are predicted to be cohabiting with a social father. This result is consistent with the result for all mothers and has a comparable level of statistical significance despite the much smaller sample size.

The second panel examines the marriage and cohabitation patterns for the 150 mothers who had no history of AFDC receipt in the two years prior to entering W-2.

Table 3 Multinomial model of living arrangements, three subgroups

	Married		Cohabiting, Father		Cohabiting, Nonfather	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
Education Less Than High School ($N = 338$)						
Experimental group	0.060	0.486	–0.168	0.410	–0.772†	0.456
Child support history in year prior entering W-2						
$1,000 or more paid on behalf of family	0.021	0.825	–1.034	0.685	0.325	0.615
Age >30 years	–1.622†	0.953	–0.887	0.702	–0.522	0.503
African American	–0.961†	0.579	–0.660	0.537	–0.996	0.512
Region (vs. Milwaukee)						
Other urban county	0.146	0.742	0.468	0.579	–1.247	0.931
Rural county	0.211	0.999	0.826	0.748	0.437	0.860
Age of youngest (vs. <3 years)						
Age 3–5 years	1.220*	0.509	0.540	0.533	0.521	0.592
Age 6 or more years	0.459	0.823	–2.670*	1.079	0.990†	0.542
Intercept	–1.665**	0.594	–1.293**	0.502	–1.327**	0.510
Log-likelihood			–289.7			
No AFDC Receipt In Prior Two Years ($N = 150$)						
Experimental group	–0.941	0.574	0.580	0.676	–2.296*	0.931
Child support history in year prior entering W-2						
$1,000 or more paid on behalf of family	–0.391	0.889	–0.101	0.984	–17.669**	0.978
Age >30 years	–0.283	1.075	–16.356**	0.756	0.784	0.833
African American	–1.020	0.744	0.085	0.979	–1.833	1.281
Education (vs. less than high school)						

Table 3 (continued)

	Married		Cohabiting, Father		Cohabiting, Nonfather	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
High school	0.939	0.593	–1.411	0.967	0.152	0.761
More than high school	–0.223	0.874	–1.167	1.410	–0.694	0.841
Region (vs. Milwaukee)						
Other urban county	–0.032	0.718	–0.277	0.975	0.043	1.150
Rural county	1.113	0.794	0.972	1.157	–0.281	1.229
Age of youngest (vs. < 3 years)						
Age 3–5 years	–17.423**	0.853	–17.278**	0.848	2.691*	1.125
Age 6 or More	–0.480	1.131	–14.361**	1.492	–0.055	1.066
Intercept	–1.054	0.910	–1.149	1.133	–0.393	1.272
Log-likelihood			–75.3			
Mothers With Youngest Child Age 5 or Younger at Entry ($N = 516$)						
Experimental group	–0.178	0.330	–0.085	0.346	–0.950*	0.381
Child support history in year prior entering W-2						
$1,000 or more paid on behalf of family	–0.297	0.563	–0.880	0.542	–0.220	0.536
Age >30 years	–0.282	0.507	–0.950	0.637	–1.170*	0.595
African American	–1.363**	0.373	–0.983*	0.446	–0.834†	0.449
Education (vs. less than high school)						
High school	0.180	0.345	–1.219**	0.434	–0.166	0.410
More than high school	0.560	0.528	–0.587	0.689	1.277**	0.467

Table 3 (continued)

	Married		Cohabiting, Father		Cohabiting, Nonfather	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
Region (vs. Milwaukee)						
Other urban county	–0.081	0.436	0.213	0.511	–0.017	0.510
Rural county	0.453	0.517	0.718	0.616	1.017†	0.597
Intercept	–0.986*	0.395	–0.904*	0.444	–1.217**	0.467
Log-likelihood			–436.5			

Notes: Models also control for period entering W-2, a proxy for long-term receipt; there are no statistically significant estimates. A Hausman-McFadden test suggests the independence of irrelevant alternatives (IIA) assumption has not been violated.

$^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$

Similar to the overall results and the results for those with low education, the estimated difference in the likelihood of cohabitation with nonfathers is statistically significant. Moreover, after we transform the coefficients, we see that experimental-group mothers without a history of AFDC receipt were more likely to be single (unpartnered) than their counterparts in the control group (95 % compared with 86 % for the control group). The relatively large effects among mothers without a history of AFDC receipt are consistent with findings from the main evaluation for other outcomes. Nonetheless, the point estimates should be interpreted with caution given small sample sizes. Although the sign and statistical significance of the effects are robust, the point estimates are sensitive to alternative specifications.

Finally, we also examine those whose youngest child was younger than age 6 at the time of entry to W-2, a group whose changes in partnering may be particularly consequential for their children. Similar to the other results, within mothers with preschool children, the women in the experimental group are less likely to be cohabiting with a nonfather several years later ($p < .02$). In all three subgroups, the control variables tend to show similar results.

In summary, when we compare mothers in the experimental and control groups, we find that mothers eligible to receive more child support are less likely to cohabit with men who are not fathers of any of the mothers' children. This result holds across subgroups of interest. In the sample as a whole, and in the subgroup of mothers with no recent AFDC history at the time of entry to W-2, those in the experimental group were more likely to be single. There are no significant differences in the probability of being married in the sample as a whole nor in any of the subgroups tested.

Discussion

In the 35 years since Hannan et al. (1977) used evidence from a random-assignment experiment to argue that cash benefits provided to low-income families led to increases in marital dissolution, a substantial literature has addressed the relationship between economic resources and romantic partnerships. The demographic context has changed substantially in the intervening years (Cherlin 2009). Nonmarital births now constitute more than 40 % of all births (Hamilton et al. 2011), and a higher percentage of births are to low-income women, highlighting the importance of understanding partnering transitions for single mothers. Growth in cohabitation (Kennedy and Bumpass 2008) and multiple-partner fertility (e.g., Cancian et al. 2011; Meyer et al. 2005) have also increased the importance of understanding partnering beyond marriage as well as partnerships with fathers versus other men.

The results presented here take advantage of a recent random-assignment experiment to evaluate the economic independence hypothesis and its relevance in the new demographic context. The results are consistent with the hypothesis that additional economic resources (in this case, increased child support) increase a mother's economic independence and reduce her need to partner in order to make ends meet: mothers in the experimental group are more likely to be single. Moreover, the increase in singleness is not offset by detectable decreases in marriage or cohabitation in general but only in the likelihood of cohabiting with a nonfather.

Although the estimated effect for the full sample is marginally statistically significant ($p < .10$), the findings regarding reduced cohabitation with social fathers among those in the experimental group are robust to alternative sets of regression adjustments and persist across all tested subgroups. As was found for many other outcomes in the original evaluation (Meyer and Cancian 2001), the effect (i.e., reduced cohabitation in the experimental group) is particularly strong among families new to the welfare system. These families may have been more responsive to the new policy because of less exposure to the previous system or because of other differences among new and continuing or returning participants.

The effects are particularly notable given the limited difference in the treatment received by those in the experimental and control groups. Those in the experimental group received all current child support paid, regardless of their W-2 participation status. Those in the control group received only part of the child support paid on their behalf, beginning with their entry to W-2 (in 1997 or 1998) until they stopped receiving TANF, or until the policy changed in 2002. Thus, the experimental assignment directly affected receipt of child support for a fairly short period for most participants. Moreover, the policy change affected the incomes of only those for whom child support was actually paid—about one-half of the mothers. Finally, the magnitude of the increase in support is modest: mothers in the experimental group were about three percentage points more likely to receive support in the first year and received 23 % more support, but the effect declined over time (Cancian et al. 2008). For all these reasons, the estimated effects of the policy on family structure are likely to be conservative.

The power of random assignment gives us confidence that we have identified a difference in partnering behavior caused by the policy difference that resulted in more child support for one group than the other. Earlier, we noted that an experiment is well suited for identifying a total effect but not the precise pathways. Random assignment assures that the difference in outcomes is not due to preexisting or unmeasured differences between the two groups, so it is a very powerful design. However, studies using random assignment are less strong at identifying how the intervention changed behavior. As a result, the change in partnering could be the result of some indirect effects of the policy rather than just a direct effect of the change in income. For example, fathers who know that the child support that they pay directly benefits their children may be encouraged to take more responsibility for their children in other ways, and the increased connections between fathers and their children could then lead to more connections between fathers and the mothers of their children. It is possible that increased connections could affect mothers' partnering behavior—not as a result of the increased income *per se* but because the policy change caused changes in fathers' attitudes. Another possibility is that when the mother receives the full amount of child support paid by the father, parents may have less conflict over whether the father is fulfilling his responsibilities. Less conflict between the parents could make it more likely for the parents to live together and less likely for either parent to form a new romantic partnership; alternatively, the lack of a conflictual previous relationship may make both parents more attractive partners for others, perhaps increasing the likelihood of their marrying or cohabiting with a new partner. Again, if the new policy causes less conflict and the lowered conflict affects partnering, the experimental group could have different partnering behavior from the control group, but it would not necessarily be the result of more income *per se*, but instead due to an indirect effect of the policy change.

One way to explore the evidence for these alternate explanations is to see if there are differences between the experimental group and the control group in the intermediate steps in the hypothesized causal chain. This is not a perfect test because even if there are differences, it would still be possible that the intermediate steps were themselves the result of additional child support income. The main evaluation (Meyer and Cancian 2001) found no overall difference between the experimental and control groups in the number of days of contact between nonresident fathers and a focal child, the extent to which fathers looked after the focal child so that the mother could go to work or school or look for work, or the extent to which mothers reported that the father was doing a good job as a parent (all $p > 0.20$). On all these intermediate variables, there was no effect in either the first year or the second year of the experiment. In the first year, but not the second, there was a small difference ($p = 0.09$) in that experimental-group mothers were less likely to report high conflict on at least one child-rearing issue (Meyer and Cancian 2001). The mixed findings on conflict hint that the experimental group's lower likelihood of living with a social father (and increased likelihood of remaining single) could be due to the difference in treatment of child support leading to less conflict. However, assessing all these results together, the lack of difference between the experimental and control groups in the levels of contact and relationship is consistent with the interpretation that the difference in partnering behavior is due to the income difference rather than another aspect of the parents' relationship. Notwithstanding this overall assessment, additional research would be useful.

The effect of economic resources on family structure has been a central focus for researchers across multiple disciplines, and related concerns have motivated significant policy interventions over that last two decades. However, relatively little research has been able to evaluate the causal effect of income on partnering. In addition, much of the research has not focused on economically disadvantaged mothers, who are particularly likely to be constrained by economic factors and who are subject to the most intense policy interventions. Finally, few studies have differentiated types of cohabitation, a distinction that has been shown to be potentially important (Cherlin and Fomby 2004) and that is increasingly relevant given demographic trends. This article, taking advantage of a unique random-assignment experiment, allows us to address some of these limitations. The results point to the importance of new understandings of family structure and new directions for research. If we were to measure only the three traditional categories of partnering status—married, cohabiting, or single—we would fail to capture the full effects of increased resources. Only when the distinction is made between cohabitation with a biological father and cohabitation with a social father do we see a more precise effect of increased resources and a way in which they enable mothers to be more independent. This empirical support for an independence effect is relatively rare (e.g., Burstein 2007) and suggests that research on the effect of resources in the new demographic context is important.

Additional research is needed to test these findings further and address the limitations of this analysis. Although low-income families are of particular interest to policymakers, the question of the relationship between income and family formation is of broad importance and requires analysis of potential differential effects across the distribution of income. Similarly, additional research identifying the causal effects of child support income (for resident parents) and child support payment obligations (for nonresident parents) on partnering is increasingly important given the growth in scope and effectiveness of the child support enforcement system for middle-income families. Moreover, while the experimental evaluation of this child support policy provides an unusual opportunity to evaluate the impact of an exogenous change in income on partnering, it is possible that changes in child support income, or child support pass-through policy, have different effects than would be generated by a change in another income source. More research to evaluate potential differential effects would clearly be useful.

The well-being of children living in low-income families is a central concern of social policy. Income support policies for families with children have undergone a remarkable transition (Moffitt and Scholz 2010). The last decade has seen a substantial expansion in public programs explicitly designed to foster marriage (Wood et al. 2012), and recent initiatives more closely integrate efforts to improve earnings, child support outcomes, and relationship quality (U.S. Department of Health and Human Services, Office of Child Support Enforcement 2012; Wood et al. 2012; Zaveri and Hershey 2010). These policy initiatives further highlight the importance of understanding the role of economic resources in shaping family formation decisions and related outcomes. Our key finding is that increased resources reduce the likelihood of living with a social father. This is consistent with the theory suggesting that disadvantaged mothers who gain extra resources are less likely to need to partner with a new man for purely economic reasons, which may have positive consequences for children's well-being (Cherlin 2009). Importantly, in this research, the increased independence does not show itself in decreased marriage, lending support to the idea that some policies can increase economic support for vulnerable families without discouraging marriage.

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