THE LABOUR MARKET CONSEQUENCES OF TEENAGE CHILDBEARING*

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Abstract

This paper provides estimates of the impact of an unanticipated child during adolescence on labour supply and earnings using data for women who gave birth between 1976 and 2015 drawn from the 1990 and 2000 censuses and the American Community Surveys. Twins at first birth are used as an instrument to avoid the problems of fertility endogeneity. Estimates from our IV models indicate that the arrival of a second-born twin had severe economic consequences for adolescent women over most of our data.

Key words**:** Teenage motherhood, fertility, female employment, instrumental variables

JEL classifications: J13, J31, J16

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1. Introduction

Since the early 1980s, teenage childbearing has been considered a matter of public concern in the United States. Despite the fact that the rate of teenage childbearing has dramatically declined in recent years, today births to adolescents in the United States account for approximately one fifth of all annual births, which is far higher than in any other developed country (Kearney and Levine 2012, UNICEF 2013)[1]. Interest in teenage childbearing has been fostered in part by a large body of cross-sectional evidence which indicates that giving birth while still a teenager is associated with a range of problems for young women including reduced labour market participation and diminished earnings (Card 1981; Hofferth 1987; Upchurch and McCarthy 1990).

Among the competing explanations for why childbearing is negatively associated with the labour market decisions of young women, four receive the most attention. The first is the curtailment of formal schooling when women are adolescents for childcare which may damage subsequent career opportunities. The second is to the extent that mothers spend more time outside the labour market for child care, the accumulation of less human capital through employment experience and on-the-job training in the post-school stages of the lifecycle will result in lower future wages. The third is the necessary reduction in the amount of effort available for market work due to effort expended on child care activities, which would reduce the earnings of women bearing greater responsibility for child care (Becker 1985). The fourth is the existence of unobservable characteristics that are correlated with both the decision to have children during adolescence and labour market outcomes. As many social scientists have argued, those women who chose to have children when they were teenagers may well be the same women whose economic outcomes would have been low in any case (Geronimus

[1] Figure 1 plots the teen birth rate from 1980 to 2016.

and Korenman 1993; Hoffman et al. 1993). The primary difference between the first three explanations and the fourth is that the first three suggest that having a child while still a teenager has a causal effect on future labour market outcomes, while the fourth implies that the estimated effects are the result of omitted-variable bias. This paper attempts to identify the causal effect of a second child in adolescence on women's labour market outcomes.

Credible estimates of the consequences of teenage births are difficult to establish. In the existing literature, a variety of econometric strategies have been used to tackle the challenge posed by the endogeneity of fertility. The most common approach is to control for a wide set of observable factors that take account of the disadvantaged backgrounds of teenage mothers when they were growing up (see, for example, Duncan and Hoffman 1990; Furstenberg 1991; Hayes 1987). These studies generally find some evidence of adverse consequences as a result of a teenage birth. For this approach to establish the true effect of teenage childbearing, it is required that the timing of fertility is uncorrelated with all unobservable factors that might influence the outcome under consideration, conditional on measured background characteristics. A second approach is based on panel data which compares the outcomes of sisters using differences in the timing of births. These studies generally provide support for the hypothesis that the effects of teenage childbearing are greatly reduced but not eliminated in models that control for family-specific omitted variables (Geronimus and Korenman 1992; Hoffman et al. 1993; Ribar 1999). Establishing causal estimates using this approach is challenging because parental inputs and childrearing practices may respond to differences in outcomes across sisters (Rosenzweig and Wolpin 1995).

This leads to a third remedy to the problem of omitted variable bias which involves the use of instrument variables to obtain consistent estimates[2]. In this study, we use an exogenous fertility event, first introduced by Rosenzwieg and Wolpin (1980), that relies on the occurrence of twins at first birth as an unplanned event to provide estimates of the effect to which teenage women's labour supply and earnings respond to an unplanned second child. One obvious drawback of using the "twins-first" approach is that while it helps to identify the marginal effect of an additional child given the presence of one child, it does not allow us to look at the impact of having one child as opposed to none as a teenager. Nonetheless, estimating the marginal effect of an additional child for teenage women is relevant and important since approximately 20 percent of teenagers who become mothers will go on to have a second child before age twenty[3].

This work is closely related to a small number of studies that use the twins-first methodology to estimate the causal effect of family size on women's socioeconomic status (Bronars and Grogger 1994; Grogger and Bronars 1993; Jacobsen et al. 1999; Vere 2011). Bronars and Grogger (1994), using data from the 1970 and 1980 USA population censuses, examined the effect of single motherhood on a range of outcomes but they did not examine the impact of teenage motherhood. Jacobsen et al. (1999) expanded their work using the same data to look in detail at married mothers, while Vere (2011) provided more recent evidence for both married and divorced women in the 1980, 1990 and 2000 censuses. Closest to the present study is Grogger and Bronars (1993) who using data from the 1970 and 1980 censuses examined the effect of teenage motherhood on a range of outcomes including marital status,

[2] Examples of instruments used to produce exogenous variation in the timing of births include abortion legislation (e.g., Angrist and Evans 1996; Ashcraft et al. 2013; Klepinger et al. 1999; Ribar 1994), miscarriage (e.g., Ermisch and Pevalin 2003; Goodman et al. 2004; Hotz et al. 2005), contraceptive legislation (e.g., Bailey 2006), and infertility (e.g., Aguero and Marks 2008).

[3] Data in the decennial censuses and American Community Surveys between 1990 and 2016 indicate that one fifth of women who had one teenage birth went on to have a second before the age of 20 (author's calculations).

labour force participation, family earnings, household income, poverty status and welfare recipiency. Their primary interest, however, was in the consequences of teenage motherhood for family instability and other joint family outcomes; they did not explore in detail the separate effects of an unplanned child on the mother's own labour market behaviour. Thus how women's labour supply and earnings responds to the arrival of an unplanned second child due to the occurrence of twins during adolescence remains largely unexplored with the twins-based instrumental variables method.

Although the United States still stands out as having one of the highest rates of teenage childbearing among developed countries, the number of births to teenage women has steadily declined over the past quarter century, falling from a peak of 61.8 per thousand in 1991. This trend can be seen in figure 1, which plots the sharp decrease in the number of births to women between the ages of 15 and 19 (Martin et al. 2018)[4]. Following the recent recession, teenage women experienced one of the largest declines in birth rates, which dropped from 41.5 per thousand in 2008 to 20.3 per thousand in 2016[5]. Given this backdrop of a decline in the prevalence of teenage childbearing, this paper investigates the impact of teenage fertility on women's labour supply and earnings. We trace women who had children born between 1976 and 2015 by looking at mothers surveyed in the U.S. censuses from 1990 and 2000 and the American Community Surveys (ACS) spanning the period from 2001 to 2016. In doing so, this research is the first to examine the impact of changes in exogenous fertility on labour supply and earnings for children born during and following the Great Recession.

[4] The causes of teenage childbearing have been recently studied by Kearney and Levine (2015, Kearney and Levine 2014).
[5] The trends in the teen birth rate display in figure 1 are also apparent in the overall population as depicted in figure 2.

Unlike earlier work, we investigate the effect of an unexpected second child on the mother's labour market status using four key measures including employment status in the year prior to the survey, weeks worked per year, usual hours per week, and own earned income. Except for employment participation, these labour market outcomes have not been previously utilized in estimating the effects of an unplanned second child on teenage women. In order to achieve a more complete picture of how the impact of a second child has affected the standard of living of teenage mothers, we also analyse family income. By estimating the effects of a second child on a broader range of outcomes than has previously been considered, the results of this article are more comprehensive in part because they are consistent across alternative measures of economic wellbeing.

A related public policy question of interest is whether postponing births can improve outcomes for young mothers. To provide some insight into this question we compare our results for teenage mothers to the results for young-adult mothers by also examining the impact of a second child for women whose first birth occurred somewhere between age 20 and 24, rather than the more extensive definition of all adult women that is usually employed.

From a policy perspective, knowing how young women's labour market decisions respond to the arrival of an unexpected child is important. If the results of this research show that an additional child substantially impedes mothers' labour force participation and earnings, then public policies with the goal of strengthening the labor force attachment of women with children would be appropriate. Such policy options may including subsidizing child care costs as labour market decisions are likely to be significantly affected by the costs associated with replacing maternal care with nonmaternal care. In addition, if the results of this research show that teenage mothers fare much worse in the labour market compared to those women who had their first child in their early twenties then policy measures designed to delay childbearing until adulthood may be desirable. Such policies may involve measures designed to reduce exposure to pregnancy including encouraging schools to provide education in reproduction, and special adolescent clinics associated with schools where young women can receive contraceptive services and counselling.

This paper is organized as follows. The next section presents our econometric model and discusses the twins-first instrument. Section 3 describes the data. Section 4 presents the results including an analysis of racial differences. Section 5 provides the conclusion.

2. Empirical Strategy

The objective of this paper is to distinguish the true causal effect of an unplanned child on women's labour market outcomes. If the occurrence of twins at first birth were truly a random event, then simple differences in the average level of the outcome variables of interest between women who had twins and those who gave birth to singletons would yield consistent estimates of the effects of exogenous variation in fertility. However, twinning probabilities are known to be biologically related to variation in mother's age at first birth and race. Failure to control for these factors in estimation could result in inconsistent estimates. To allow for this potential problem, it is necessary to estimate the following IV model for the sample of young women with at least one child:

$$y_i = \beta_0 + \beta_1 n_i + \beta_2 A_i + \beta_3 X_i + \mu_i \qquad (1)$$

$$n_i = \alpha_0 + \alpha_1 twins_i + \alpha_2 A_i + \alpha_3 X_i + \varepsilon_i \qquad (2)$$

Equation (2) represents the first-stage of the IV estimation, where equation (1) is the second stage. The dependent variable y_i is a measure of the outcome variable (e.g. employment participation, earned income) [6]. The n_i denotes the total number of children born to individual i by the survey date. The instrumental variable is $twins_i$ which is a dummy variable set to one if a teenage mother had two children at the first birth. The A_i represents mother's age (measured in quarters) at the time of her first birth. The other exogenous control variables in the vector X include a quadratic for mother's age at first birth, a quadratic for time since the first birth, survey year dummies, state of birth dummies, state/year unemployment rate and a series of controls for race, ethnicity and marital status[7].

In addition to IV estimates, in the appendix to this paper, we also provide OLS estimates of the relationship between women's outcomes and family size for comparison with much of the existing literature. In general, the OLS regressions only yield consistent estimates of increments to family size if there is no correlation between the error term and family size in equation (1), conditional on the set of control variables. OLS estimates will be biased if there is a correlation between family size and some characteristic (e.g. career ambition) excluded from the control vector that also affects labour market outcomes. If this omitted variable were positively correlated with labour market success and negatively correlated with the number of children, excluding it from the OLS regression would bias estimates upwards (in absolute

[6] The summary statistics displayed in tables 1A and 2A for teenage mothers and young-adult mothers respectively show that working for pay falls for women with twins in both groups. If the types of mothers who drop out of the labour force are positively or negatively selected, this will bias the earnings estimates. Using the census and ACS data, we are unable to include a correction for selection in our earnings regressions. However, a comparison of the demographic variables presented in tables 1A and 2A for those working for pay who have twins and those who have singletons with the overall samples of each group reveal no statistically significant differences. Unsurprisingly, marriage rates are marginally lower and family income marginally higher among those working for pay in both groups, though these differences are also not statistically significant. Therefore, it appears unlikely that there are important labour force selection differences between women who have twins and women who have singletons at first birth.

[7] The variable "time since the first birth" is calculated as mother's age (in quarters) at the time of the survey minus her age (in quarters) at the time of her first birth. State/year unemployment rate is from the Bureau of Labor Statistics. Race is grouped by three dummy variables: white, black and other races (which includes American Indian or Alaska Native, Chinese, Japanese, other Asian or Pacific Islander, other race (not elsewhere classified) or multiple races). In all specifications, an additional dummy variable was included that controls for Hispanic ethnicity. We control for race since the fraction of twins is higher among blacks than whites (Myrianthopoulos 1970).

terms). The IV approach can be used to obtain consistent estimates of β_1 if it can be successfully argued that a twin birth is exogenous to the error term and sufficiently correlated with family size.

Although many researchers have regarded the twins instrument as an ideal instrument with which to identify the causal effect of changes in fertility on women's labour supply and earnings, there are a few important caveats. First, two children arriving at the same time may have adverse effects for the mother and the child in terms of their health. These implications could have direct effects on the mother's ability to work, which would violate the exclusion restriction. To the extent that twins have a direct negative effect on the labour market behaviour of the mother, our estimates of family size using the twins instrument will be biased towards finding larger negative effects. However, the opposite bias may also arise; with zero spacing between twins, there could be opportunities for economics of scale in the provision of resources including the mother's time for childcare which could reduce the trade-off between childcare and paid work. It is not possible to empirically examine these issues using the data at hand.

A second concern is that abortion can make the occurrence of twins non-random. This might arise if teenagers who choose to abort twins are those who have better labour market prospects (Ashcraft et al. 2013). This effect may compromise the validity of the twins instrument. To the extent that abortion is more frequent among black women, it is possible that the instrument is less valid for this subsample of teenagers in particular (Angrist and Evans 1996). Our expectation is that ignoring this selection issue leads to a downward bias (in absolute terms) in estimates of the effect of an additional child on labour market outcomes. Our IV estimates represent a local average treatment effect for those women who choose not to abort and go on to carry the children to term.

Finally, consistent with previous research, the summary statistics in tables 1A and 2A show that blacks are overrepresented in our samples of twins[8]. Our fertility estimates for blacks in the 1990 census may violate the exclusion restriction given that the black sample could include biracial blacks who are less likely to give birth to twins and may face less discrimination in the labour market than monoracial blacks (Fairlie 2009). However, the 2000 census and the ACS surveys allow individuals to report more than one race. In analysis not reported, as a check we included the two race identifier in regressions using these data. This had very little effect on our estimates.

3. Data

Our data are drawn from the 1 percent and 5 percent samples for the 1990 and 2000 U.S. decennial censuses and the ACS between 2001 and 2016 accessed from the Integrated Public Use Microdata Series (IPUMS). In order to examine the effects of fertility on women's labour supply and earnings over time, we pool successive cross sections of data from the decennial censuses and the ACS. The decennial censuses and the combined ACS data surveys create a large sample for the relatively small target population of teenage women who experienced twins at first birth.

The data record the number of own children living in the household for each individual, though it does not record the number of children ever born to each female respondent (with

[8] Previous research by Bronars and Grogger (1994), Jacobsen et al. (1999) and Vere (2011) also found that blacks are more likely to give birth to twins than whites.

the exception of the 1990 census where both variables are available). To extract mothers living with their biological children, we first discard all observations where a step-mother is present in the household. Then using a procedure similar to that outlined by Angrist and Evans (1996, p. 30), mothers are matched to their children using the detailed relationship codes. Our analysis focuses on the subset of women who were teenagers at the time of their first birth. Thus we restrict the sample used to: (1) women who gave birth to their first child after age 14 and before age 20; (2) are between 20 and 30 years of age at the time of the survey; and (3) whose oldest child was at most 16 years of age at the time of the survey[9]. The sample excludes women over the age of 30 because it is not possible to determine conclusively whether their first-born child is the oldest child still living in the household[10]. The sample also excludes women who were born outside the US.

In studying the probability of twinning in the 1990 and 2000 censuses and in the ACS between 2001 and 2004, it is important to note that twins are identified as children who are born in the same year, as quarter of birth data is not reported in these surveys. Thus we find that 1.03 percent of all first births in these data are twins. In the ACS data between 2005 and 2016, twin births are accurately identified by the year plus the quarter of birth. In this sample, the probability of a twin first birth is 0.76 percent. Our figures can be compared with the 0.0069 probability of twinning found by Grogger and Bronars (1993) for a sample of teenage mothers from the 1980 census in which quarter of birth is recorded and used to identify twins. The jump in the probability of twinning over time is consistent with vital statistics data

[9] Unlike the 2000 census and the ACS surveys, in the 1990 Census respondents were asked the number of children ever born as well as the number own children living in the household. Using these variables along with the sample restrictions applied in this paper we estimate that 9.5% of mothers do not live with either one or more of their biological children. This is a higher percentage of missing children than in the wider population (which we estimate at 6%). We are unable to extract from the data where in the order of birth the missing child is placed and have no way of knowing whether or not it is the first born child.

[10] The sample also excludes women whose oldest child is 9 months or younger at the time of the survey. This restriction is used because labour market outcomes in the ACS are reported for the year prior to the survey.

tabulated from birth certificates in the United States which shows that across all births the twinning rate rose from 1980 through 2016 (Martin et al. 2018). The rise in multiple birth rates has been associated with the adaptation of bovine growth hormones in the early 1990s (Steinman 2006). Since almost everyone consumes dairy products, this factor should not lead to biases in the twins-based IV estimates.

The use of year of birth to identify twins naturally generates a much higher estimate of the number of twins since two children born in the same twelve months period are classified as twins. This is a limitation of using the censuses and ACS data prior to 2005. Vere (2011) points out that any measurement error arising from using year of birth data alone to identify twins is only problematic when it is correlated with the error term in the second stage equation (equation 1 above). In unreported analysis, to determine if such a correlation exists over-identification tests were conducted with the ACS 2005-2016 data which contains quarter of birth in each year. The results from these tests suggest that there is no evidence that the IV estimates of the effects of an additional child on any of the dependent variables are different depending on whether quarter of birth data along with year of birth is used to generate the instrument.

It is useful to compare our results for women who had their first child before age 20 with those who had their first birth between age 20 and 24[11]. This sample contains a higher fraction of twins-first mothers, which reflects the well-known biological relationship between

[11] For this analysis, the sample includes women who had their first child between 20 and 23 ¾. The sample is cut off at age 23 ¾ to avoid any selection arising from fertility treatments (such as *in vitro* fertilization) among older women which is known to increase the likelihood of twins. This sample is restricted to women between ages 24 and 30 at the time of the survey whose eldest child is no older than age 10. In robustness analysis (not reported), we also explored the sensitivity of our results for teenage mothers to imposing the same restriction in terms of the child's age. That is we limited the teenage mother sample to include only mothers whose oldest child is no older than 10. These estimates, which are noisier, are similar to the estimates reported below in table 5.

older age at conception and the greater probability of giving birth to twins (Waterhouse 1950). Means and standard deviations of these and other demographic and socioeconomic variables used in this study are displayed in tables 1 and 2 for the overall samples of teenage mothers and young adult mothers, respectively.

Given the entire length of our data, we have gathered retrospective information on birth histories for adolescent mothers who gave birth to their first child between 1976 and 2015. Using this time series we divided the data into five eight-year sub-periods (1976-1983, 1984-1991, ...,2008-2015) according to the year that the mother gave birth to her first child. The analogous information is only available for young-adult mothers who gave birth to their first child over the period from 1984 to 2015 as a result of the restrictions imposed on our sample. Our motivation for including separate analysis by these sub-periods is to see whether the effects associated with an additional child vary with changes in the teen birth rate, changes in public policies designed to induce women with children to enter the labour market and changes in the business cycle over time. Figure 1 plots trends in the teen birth rate over time and shows that most increases in the birth rates were concentrated in the 1984-1991 sub-period with the birth rate reaching a peak in 1991[12]. A long-term decline in the teen birth rate began in 1992. This significant change in fertility coincided with changes in welfare and tax policy designed to make work more attractive for single mothers with children. Most notably, the passage of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 (PRWORA) significantly reduced the generosity of the welfare system with respect to mothers with young children[13]. Although PRWORA was adopted in October 1996, several

[12] Figure 2 shows that these trends for teenage women reflect trends for the entire population.
[13] Under PRWORA, Aid to Families with Dependent Children (AFDC), was replaced in 1997 with Temporary Assistance for Needy Families (TANF). TANF introduced lifetime time limits, usually of no more than 60 months, for receipt of cash assistance and new work-conditioned welfare benefits which generally did not exempt mothers of young children unlike AFDC. In addition, federal expenditure on child care subsidies was

states had begun serious welfare reform activities as early as 1992 similar to those that would be adopted later under PRWORA (for a review, see Grogger 2003)[14]. These changes to welfare policy were accompanied by dramatic changes to the Earned Income Tax Credit (EITC) [15]. Unlike earlier expansions, those that occurred after 1993 significantly increased the take-home earnings of very low income women especially those with two or more children. The economic incentives associated with these public policies may have dramatically increased the opportunity cost of having children and changed the composition of women entering motherhood during adolescence over the course of our data. Concomitant with these reforms and the declining teen birth rate, women's labour market participation sharply rose throughout the 1990s. Figure 1 also reveals that the most dramatic decline in the teen birth rate occurred in the years through and following the Great Recession. For the entire sample of teens, the birth rate declined by 49% between 2008 and 2016. This decline may reflect the impact of the economic recession that started in 2008[16]. The potential implication for our study is that if fewer women for whom motherhood would be costly gave birth following the recent economic downturn, this would lead to a smaller effect for a second child during this period[17].

massively increased over the 1990s (and expanded in the 2000s) owing in large part to the creation of the Child Care and Development Fund (CCDF).

[14] Grogger (2003) using data between 1978 and 1999 showed that the rate of welfare utilisation among female-headed families peaked in the early 1990s with female labour supply and earnings rapidly rising over the 1990s.

[15] EITC credits increased from $1.6 billion in 1984 to $25.1 billion in 1996. In 1991 the EITC was expanded to provide a larger credit for families with two or more children. Also from 1991, the EITC was not counted as income in most means-tested programmes increasing its value for low-income mothers.

[16] On the one hand, a weaker labour market following 2008 would make having children less affordable. However, on the other hand, worsening job opportunities would also lower the opportunity cost of having children making it more likely for women to give birth.

[17] Recent empirical evidence by Dettling and Kearney (2014) using data from 1990 to 2007 indicate a role for house prices (estimated to be the largest annual cost of raising a child) in altering the demand for children. Earlier work by Lleras-Muney (2004) found that white women are more likely to give birth during an economic downturn while black women are less likely to give birth.

The variable of primary interest in this study is the number of children. The summary statistics displayed in table 1 and table 2 show that teenage mothers tend on average to have more children than young-adult mothers. The dependent variables that are considered here include: (1) whether the mother worked during the year prior to the survey; (2) the number of hours she usually worked, if she worked in the year prior to the survey; (3) the number of weeks she worked during the year prior to the survey; (4) mother's own annual earned income in the year prior to the survey; and (5) total family income earned from all sources in the year prior to the survey. Family income is introduced alongside mothers' own labour market outcomes to measure changes in broader family well-being. The summary statistics presented in table 1 indicate that from 1976 to 2015 the growth in adolescent mothers' labour supply is substantial with most of the increase in women's labour supply and real own earnings (measured in constant 1999 dollars) having occurred for women who had their first child by the end of the 1990s[18].

Following the relatively mild recession of the early 2000s, labour supply and women's own earnings began to fall, though labour supply continued to remain well above its 1976-1983 level. However, for the most recent cohorts of adolescent mothers, own labour market earnings are lower than for all earlier cohorts. Also we observe that successive cohorts of adolescent mothers have experienced steady declines in real family income for the entire period under consideration. It is clear that increases in women's participation and work effort had not been large enough to fully offset declines elsewhere in family income.

[18] Two studies examined the effect of welfare policies introduced in the 1980s and 1990s for employment and earnings among low-income women. Grogger (2003) found that the Earned Income Tax Credit was among the most important policy measures for explaining the rise in labour supply and earnings among female headed families between 1993 and 1999. Meyer and Rosenbaum (2001) showed that a large share of the increase in employment of single mothers between 1984 and 1996 could be attributed to the expansion of the Earned Income Tax Credit and other tax changes, with smaller shares for welfare benefit cuts, welfare waivers, and child care programs.

These changes in young women's labour market behaviour have occurred simultaneously with a steady increase in the share of teenage mothers who are Hispanic or in the "other race" categories. The percentage of teenage mothers who are Hispanic has grown from 8 percent in 1976-1983 to 18 percent in 2008-2015. This trend is largely attributable to a growing Hispanic population: the proportion of the teenage population that is Hispanic has more than doubled over this period. The share of black teenage mothers has not remarkably changed. The other important factor is the sharp reduction in marriage rates of teenage mothers. The summary statistics show that the marriage rate dropped 17 percentage points from 69 percent in 1976-1983 to 52 percent by 2008-2015.

The trends for teenage mothers are very close to the trends observed for young adult mothers presented in table 2. At each point in time, teenage and young-adult mothers appear equally likely to work and have similar hours and weeks worked, but average real mothers' earnings and average family earnings are remarkably higher for young-adult mothers than for teenage mothers. The parallel trends in labour force behaviour and family income for these two groups indicate that the forces impacting mothers with children are somewhat the result of changes affecting the entire female labour market, though their impact is much more pronounced among women who had their first child as a teenager. While we do not know what causes teenage mothers to experience lower family earnings than young adult mothers, it may reflect differences in spousal presence and earnings as well as income from other sources including differences in welfare income. In addition, the demographic patterns for young-adult mothers are similar to those experienced by teenage mothers though young-adult mothers are far more likely to be married and have considerably greater family income.

There are a number of reasons why we must be cautious about drawing strong policy conclusions from comparisons between teenage mothers and young-adult mothers. The counterfactual outcomes of women who gave birth during adolescence, should they have postponed childbearing until adulthood, might be less favourable than those who actually had children when they were in their early twenties. This makes it difficult to gauge the impact of delaying childbirth until adulthood. Notably, the educational outcomes that adolescent mothers would have obtained had they not given birth may be quite different from the actual outcomes of young-adult mothers that we can observe in our data[19]. Also, family background differences between women who have first births as teenagers and those who have first births in their early twenties may substantially differ. Women who give birth during adolescence are potentially drawn from more disadvantaged backgrounds than those who delay childbearing until sometime later. Thus, the estimates derived from young-adult mothers are likely to overstate the true impact of delaying parenthood until one's early twenties for women who had children during adolescence. Although our data contains no information on socioeconomic family background, a comparison of tables 1 and 2 reveals some discrepancies in demographic characteristics between teenage mothers and young-adult mothers. The tabulations show that teenage mothers are more likely to be black and Hispanic. Differences in marriage probabilities between teenage mothers and young-adult mothers indicate that teenage mothers are much less likely to be married at the time of the survey. Indeed, if young-adult mothers were more likely to be married at the time of the birth of their first child, this may suggest that they were planning on getting pregnant and did not anticipate births to be costly for labour market outcomes. In this case, estimates of the impact of

[19] Previous research by Lang and Weinstein (2015), using data from five waves of the National Survey of Family Growth on teenage childbearing in the 1950s and 1960s, found that women who gave birth before age 18 were less likely to complete 12th grade prior to Roe v. Wade. Ashcraft et al. (2013), using data from the 1995 wave of the National Survey of Family Growth, found that average education is lower by about 0.15 years as a result of a first birth before age 18, though the probability of obtaining a high school diploma was unaffected.

delaying childbearing extracted from this sub-sample would be biased towards a conservative view of delaying births until early adulthood. These important caveats should be borne in mind when drawing comparisons between the two groups in our analysis.

4. Results

4.1 The first-stage results

Table 3 presents the first-stage results that show the effect of twins in the first birth on the number of children for mothers who had their first birth as a teenager over the period from 1976 to 2015. On average teenage mothers with twin-first births have between 0.662 (SE = 0.043) and 0.855 (SE = 0.043) more children than teenage mothers with a single infant, depending on sample sub-period. The increment to family size is generally larger for blacks than whites, though these differences are never statistically significant in any sample period. The associated *F*-statistics range from 42 to 1075, which indicate that there is ample explanatory power in the first-stage regressions (Staiger and Stock 1997). Across all sub-periods the smallest partial R-squared on the instrument is 0.003 which compares favourably with those reported in Bound et al. (1995).

The analogous results for mothers who had their first child between 20 and 24 years old are shown in table 4. For young adult mothers, the effect of the occurrence of twins at first birth is to increase family size between 0.724 (SE = 0.026) and 0.772 (SE = 0.023) children. The *F*-statistics for the first-stage effects lie between 112 and 3,438 while the smallest partial R-squared is 0.010. These results suggest that the instrument is strongly correlated with the endogenous regression and therefore our IV estimates are unlikely to suffer from bias due to weak instruments[20].

4.2 The IV results

Table 5 contains the IV estimates of the effect of an additional child on subsequent socioeconomic outcomes for a sample of women who had their first birth while they were still teenagers[21]. In reading the results of these tables it is important to bear in mind that they are measuring labour supply behaviour and earned income during the calendar year preceding the survey. The coefficients in these tables are to be interpreted as the additional effect of having had a second child at first birth over and above the effect of having a single child. OLS estimates are presented in tables 3A in the appendix to this article. In general, the IV estimates imply a somewhat smaller (in absolute terms) causal effect of additional childbearing on our measures of economic well-being than the OLS estimates. This is intuitive in that we expect family size to be negatively correlated with unobserved productivity.

The estimates across all indicators of economic status for the full 1976-2015 sample period tell a fairly consistent story: labour supply, work effort and earned income fall following the arrival of a second child. Specifically, the IV results over the whole period for the entire sample of teenage mothers indicate that an additional child lowers the probability of working for pay by 6.9 percentage points. To put the magnitude of the effect in perspective, since 73.4 percent of women are in employment, the coefficient implies that an unanticipated second

[20] Staiger and Stock (1997) suggest that if the first-stage *F*-statistic is less than ten it would raise concerns that the instrument were weak.
[21]We also carried out this analysis using probit in the case of the dummy dependent variable, and Tobit for the four continuous dependent variables. These results, which do not differ qualitatively from those presented below, are available upon request.

child lowers the employment participation rate by 9.4 percent which is somewhat smaller than the results found by others in the literature (Carrasco 2001; Vere 2011)[22]. Our results also show that a second child leads to a reduction in time spent at work of 2.42 hours per week and 2.65 weeks per year. As one would expect, reductions in employment and the decline in hours and weeks of work associated with an additional child translate into lower earnings. The coefficient indicates that an additional child causes women's own earned income to fall by $991 (1999 constant dollars), which is statistically significant at all conventional levels of significance. Considering that the average earned income among teenage women in the entire sample was $9,750, this amounts to about 10 percent of the overall sample mean. The results show that the loss in family income associated with a second child is virtually identical in magnitude to that reported for women's own earned income, though these losses are much less statistically significant relative to the effects on mother's own earned income[23]. Our results can be compared with those of Grogger and Bronars (1993) who using a sample of teenage mothers drawn from the 1980 census estimated that on average family earnings fell by $1,831dollars (1999 constant dollars) for women who had a twin first birth compared to those who had a singleton first birth. Looking at our results disaggregated by race, across all indicators of labour market behaviour and income the family size coefficients for white and black women closely resemble in magnitude (and are never statistically significantly different from) those reported for the aggregated sample.

[22] Vere (2011) using the 1990 census found that a second child reduces labour force participation by 13.8 percent while Carrasco (2001) using 1986-1989 data from the US Panel Study of Income Dynamics found an average causal effect of an additional child on female labour force participation of -12.9 percent.

[23] It is worth noting that for 1992-1999, the coefficients on own earned income and family income are $1,429(SE = 387) and $1,524 (SE = 615) respectively which are quantitatively much larger than for the other sub-periods, though never statistically significantly different. Changes in tax and welfare policies as well as the general rise in female labour force participation over this period may underlie the fact that these estimates are larger than for the other sub-periods.

The subsequent columns of the table focus on five subsamples of mothers who gave birth at different times over the past 40 years. The reported coefficients, taking into account standard errors, are for the most part similar in magnitude and statistically indistinguishable to the pooled sample of observations for the periods 1976-1983, 1992-1999 and 2000-2007. For the sub-period 1984-1991 there appears to be some statistically significant effect on working for pay and number of hours worked, but no statistically significant effect on weeks worked and income. However, standard errors are large and the results are generally consistent with, but somewhat smaller than, the coefficients in the adjacent sub-periods. The weakened coefficients may be linked to the rise in the teen birth rate over this sub-period which peaked in 1991. The teenagers who gave birth may not have expected births to have been so costly for labour market outcomes, though the costs (which would be larger for mothers of twins) may have risen in subsequent sub-periods perhaps due to the welfare and tax policy changes outlined above that encouraged market work among women with children. In the sub-period 2008-2015, which are the years through and following the Great Recession, the point estimates are generally small and statistically insignificant, although in many cases the standard errors are imprecise enough that the estimated coefficients are often not statistically significantly different from earlier sub-periods (where results are significant)[24]. In this sub-period if the teenagers who gave birth were those who were less likely to join the labour force in any case, this would imply smaller negative effects because of a change in composition of who was giving birth during adolescence.

Are women who have children in their early twenties in a better position than those who have children in adolescence? Table 6 presents the IV estimates for women whose first child

[24] The large standard errors around some of these coefficients may be attributable to the relatively small sample size for this group. For the sub-period 2008-2015 the sample size is of 14,504 observations which is one-quarter of the size of 2000-2007 sub-sample. This is due to the sample restrictions imposed on the data.

occurred between 20 and 24[25]. In general the IV results for young-adult mothers tell the same basic story to those for teenage mothers in that the impact of an extra child is to significantly reduce labour supply and earnings. Across the pooled sample (1984-2015), one extra child lowers the probability of working by 6.0 percentage points (8.1 percent), reduces annual weeks worked by 3.53, and decreases mean hours per week by 2.69. Paralleling these results, a second child leads to a decrease in women's annual earned income of $1,776, which is about 15 percent of the overall sample mean. The drop in family income entirely mirrors the differential associated with a second child in women's own earned income. In addition, we observe no statistically significant racial differences between whites and blacks.

Looking across the remaining columns of the table which pool estimates over eight-year sub-periods, in the first three sub-periods the coefficients are similar in magnitude to those for the aggregated sample. The stability of these coefficients indicates that the cost of births for mothers of twins must have remained fairly stable over time. For the final period 2008-2015, on none of the dependent variables does the coefficient on family size have a statistically significant effect on labour market outcomes or family income. However, in all cases, the effects associated with a second child are similar in magnitude to those for the preceding sub-period, though standard errors are often more than doubled which is due to the smaller sample size for this group. These findings are in line with the experience of teenage mothers who showed some weakening in the response to a second child along all dimensions of labour supply behaviour over this period.

Comparing the results for teenage mothers with those of young-adult mothers is complicated by the fact that the aggregate samples do not entirely overlap. Thus, the coefficients in those

[25] The analogous OLS estimates are presented in table 4A of the appendix to this article.

columns are not directly comparable. Instead, we are able to make comparisons over the four sub-periods that are in common. Beginning with 1984-1991, the impact of an additional child for labour supply and earnings is generally felt more strongly for young-adult mothers than teenage mothers with highly statistically significant differences apparent in the number of weeks worked and mother's own earned income. It is particularly notable that for the entire sample of young-adult mothers, the impact of an additional child is to lower mother's earned income by $2,350 which is statistically significant from zero and roughly six times larger than the analogous coefficient presented for teenage mothers. While births to women of all ages trended upwards during this sub-period it would appear that young-adult mothers faced significantly higher costs associated with an additional child in terms of their own earned income forgone. Comparisons across the three other sub-periods, when birth rates trended downwards, reveal no statistically significant differences on any of the labour market outcomes between the two groups, though the effects are still generally larger in magnitude for young-adult mothers. Taken together, these findings suggest that in recent years there is no important difference in labour market outcomes between women pre- and post-adolescence who give birth to a second child due to twinning.

5. **Conclusion**

A striking feature of the United States labour market during the past quarter century has been the dramatic rise in the number of single mothers entering employment. An equally remarkable development that accompanied this trend has been a steady decline in the teenage birth rate. This article provides evidence on the causal effect of an unplanned child due to twins in the first birth for the years 1976 to 2015 on various measures of women's socioeconomic outcomes. Consistent with earlier studies using the twins-first instrument our results ranging over the entire length of our data generally reveal that an additional child significantly reduces labour supply and work intensity while imposing substantial costs in terms of forgone earnings (Bronars and Grogger 1994; Grogger and Bronars 1993; Jacobsen et al. 1999; Vere 2011). In our racially disaggregated samples, the results are similar for white and black women alike. In addition, a comparison of our results for adolescent mothers with those for women who postponed motherhood until their early twenties, show that the arrival of a second-born twin child has equally severe economic consequences for young-adult mothers. These findings are at odds with the widespread belief that early childbearing is more detrimental to adolescent women and the opposite of what one would predict on the basis that postponing childbearing until adulthood would improve their economic outcomes. Therefore policies that could successfully delay childbearing beyond adolescent would not necessarily be successful in attenuating the socioeconomic consequences of an additional child.

When interpreting our results, three caveats should be kept in mind. First, IV estimates do not identify the average causal effect for the whole population, but identify a Local Average Treatment Effect for the subpopulation influenced by the instrument (Ebenstein, 2009; Imbens and Angrist 1994). This means that our IV estimates represent the average causal effect in the population of women who had an unplanned child following the occurrence of twins in the first birth. It is important to note that having twins at first birth does not necessarily raise the number of children in a family by one child as a mother of twins may reduce her future childbearing to compensate for the unplanned child. As a consequence, our estimates do not apply to the population of women who planned to have only two children and have no additional children beyond the arrival of twins at first birth. Furthermore, our estimates do not necessarily measure the consequences of increased family size by other instruments[26]. Second, this study identifies the marginal effect of having two children compared to having one child. Most of the literature on the causal effect of teenage childbearing seeks to estimate the effect of having at least one child as a teenager relative to having no children as a teenager. If mothers of twins spend exactly twice the amount of time on child care as do the mothers of singletons (there are no economies of scale), the estimates identified by the twins approach would be the same as those that result from an unplanned teenage first birth. If women who already have one child are better able to adapt to the presence of another child, through economies of scale, then our IV results would provide conservative estimates of the effect of an unplanned singleton birth. As there are likely to be more opportunities for economies of scale in the rearing of twins, we expect that the effect of having only one child as a teenager is likely to be at least as large as the effect of having a second child as a consequence of twinning. Third, the socioeconomic data studied here relate to women's economic status up to 30 years old. Whether our results remain beyond age 30 is a question that we cannot address with our data.

[26] For example, Ashcraft et al. (2013), using miscarriage as an instrument and accounting for selection arising from abortion, found that the probability of working dropped by 5 percentage points following giving birth as a result of a first teen pregnancy before age 18. Conditional on working, they found that the number of weekly hours dropped by approximately 4 but there was virtually no effects on own annual earnings or family income from a teen birth.

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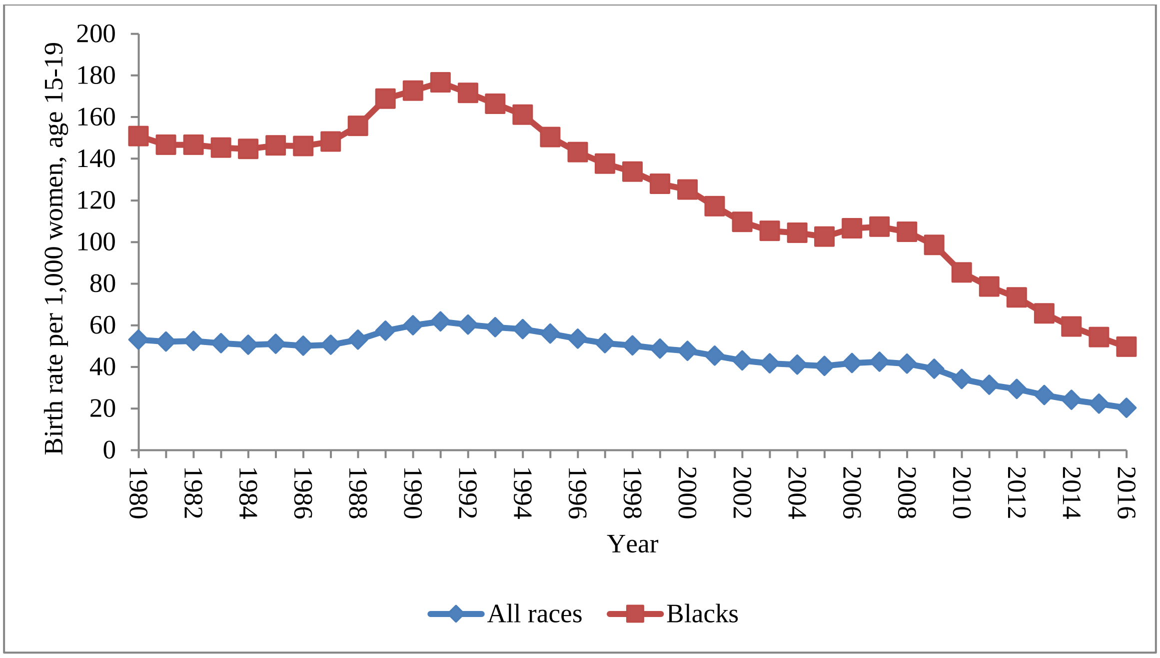
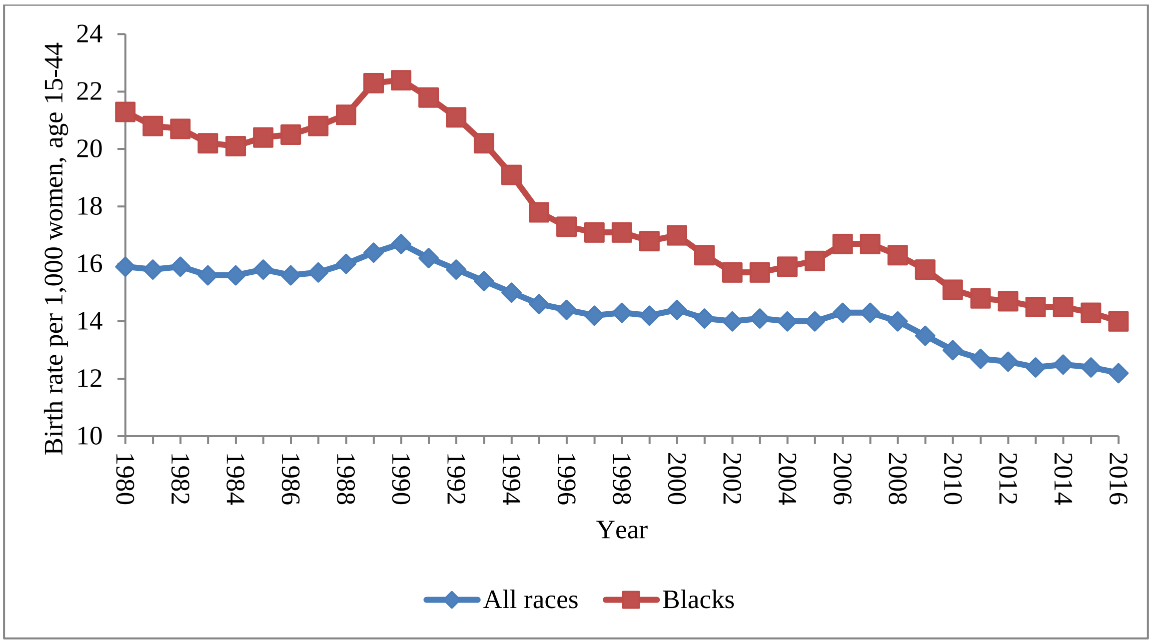
Figure 1. Trends in the teenage birth rate



Source: Martin *et al*. (2018)

Figure 2. Trends in the total birth rate



Source: Martin *et al*. (2018)

Table 1. Summary statistics by year of child’s birth, 14 ≤ mother’s age first child < 20.

	1976-2015	1976-1983	1984-1991	1992-1999	2000-2007	2008-2015
Number of children	2.217	2.435	2.163	2.210	2.169	1.712
	(1.026)	(1.037)	(1.032)	(1.011)	(1.010)	(0.803)
Twins first	0.009	0.010	0.010	0.010	0.008	0.008
	(0.095)	(0.098)	(0.099)	(0.098)	(0.088)	(0.088)
Age at 1st birth (years)	17.899	17.603	17.788	17.867	18.276	18.652
	(1.314)	(1.298)	(1.283)	(1.309)	(1.210)	(0.995)
Age at survey (years)	25.487	27.448	24.844	25.310	25.043	22.512
	(3.061)	(1.932)	(3.369)	(2.985)	(2.868)	(1.891)
Worked for pay	0.734	0.687	0.712	0.771	0.743	0.736
	(0.442)	(0.464)	(0.453)	(0.420)	(0.437)	(0.441)
No. of weeks worked	28.609	26.356	26.485	30.419	30.073	29.229
	(22.162)	(22.419)	(22.073)	(21.695)	(22.302)	(22.368)
No. of hours worked	26.120	24.667	25.703	27.728	25.764	24.589
	(18.299)	(19.249)	(18.795)	(17.661)	(17.754)	(17.501)
Mom's income (1999	9,750	9,084	9,388	10,919	9,393	7,433
dollars)	(12367)	(11357)	(13339)	(13137)	(11283)	(9407)
Family income (1999	27,532	31,648	27,043	28,326	23,945	20,032
dollars)	(24374)	(24960)	(24487)	(25037)	(22190)	(19444)
White	0.675	0.728	0.674	0.652	0.662	0.666
	(0.468)	(0.445)	(0.469)	(0.476)	(0.473)	(0.472)
Black	0.235	0.210	0.245	0.244	0.235	0.226
	(0.424)	(0.408)	(0.430)	(0.430)	(0.424)	(0.418)
Other race	0.090	0.061	0.080	0.104	0.103	0.108
	(0.286)	(0.240)	(0.271)	(0.305)	(0.305)	(0.310)
Hispanic ethnicity	0.122	0.080	0.098	0.129	0.168	0.177
	(0.327)	(0.272)	(0.297)	(0.335)	(0.374)	(0.382)
Married at time of survey	0.598	0.690	0.609	0.582	0.535	0.517
	(0.490)	(0.462)	(0.488)	(0.493)	(0.499)	(0.500)
Observations	313,921	64,779	67,634	102,968	62,709	14,504

Note: The teenage sample composition by survey year is as follows. The survey that comprises the 1976-1983 sub-sample is the 1990 census (64,779 observations). The 1984-1991 sub-period is comprised of: 37,818 observations from the 1990 census; 25,119 observations from the 2000 census; 2,006 observations from the 2001 ACS; 1,192 observations from the 2002 ACS, 742 observations from the 2003 ACS; 212 observations from the 2004 ACS; 309 observations from the 2005 ACS; 163 observations from the 2006 ACS; and 73 observations from the 2007 ACS. The 1992-1999 sub-period is comprised of: 45,463 observations from the 2000 census; 3,472 observations from the 2001 ACS; 3,531 observations from the 2002 ACS; 4,277 observations from the 2003 ACS; 4,156 observations from the 2004 ACS; 10,519 observations from the 2005 ACS; 9,265 observations from the 2006 ACS; 7,950 observations from the 2007 ACS; 5,052 observations from the 2008 ACS; 3,666 observations from the 2009 ACS; 2,680 observations from the 2010 ACS; 1,583 observations from the 2011 ACS; 683 observations from the 2012 ACS; 287 observations from the 2013 ACS; 97 observations from the 2014 ACS; and 17 observations from the 2015 ACS. The 2000-2007 sub-period is comprised of: 120 observations from the 2001 ACS; 342 observations from the 2002 ACS; 687 observations from the 2003 ACS; 939 observations from the 2004 ACS; 3,284 observations from the 2005 ACS; 4,170 observations from the 2006 ACS; 5,182 observations from the 2007 ACS; 5,236 observations from the 2008 ACS; 5,900 observations from the 2009 ACS; 6,277 observations from the 2010 ACS; 6,691 observations from the 2011 ACS; 6,306 observations from the 2012 ACS; 5,618 observations from the 2013 ACS; 4,698 observations from the 2014 ACS; 4,092 observations from the 2015 ACS; and 3,167 observations from the 2016 ACS. The 2008-2015 sub-period is comprised of: 165 observations from the 2009 ACS; 537 observations from the 2010 ACS; 1,033 observations from the 2011 ACS; 1,596 observations from the 2012 ACS; 2,026 observations from the 2013 ACS; 2,549 observations from the 2014 ACS; 3,044 observations from the 2015 ACS; and 3,554 observations from the 2016 ACS.

Table 2. Summary statistics by year of child's birth, 20 ≤ mother's age first child < 24.

	1984-2015	1984-1991	1992-1999	2000-2007	2008-2015
Number of children	1.873	1.822	1.921	1.926	1.723
	(0.822)	(0.767)	(0.836)	(0.863)	(0.785)
Twins first	0.012	0.012	0.013	0.011	0.011
	(0.107)	(0.107)	(0.111)	(0.103)	(0.103)
Age at 1st birth (years)	21.705	21.685	21.466	21.830	22.214
	(1.137)	(1.120)	(1.089)	(1.143)	(1.108)
Age at survey (years)	26.758	26.298	27.175	27.030	25.985
	(1.917)	(1.762)	(1.942)	(1.934)	(1.689)
Worked for pay	0.741	0.713	0.772	0.738	0.722
	(0.438)	(0.452)	(0.419)	(0.440)	(0.448)
No. of weeks worked	30.429	27.718	32.032	31.111	31.045
	(22.239)	(22.247)	(21.709)	(22.436)	(22.706)
No. of hours worked	26.038	25.010	27.707	25.564	24.735
	(18.274)	(18.708)	(17.899)	(18.150)	(18.178)
Mom's income (1999 dollars)	11,556	10,153	12,805	11,786	10,839
	(13515)	(12454)	(14145)	(13639)	(13532)
Family income (1999 dollars)	36,067	37,283	36,848	34,932	33,083
	(28008)	(26554)	(28836)	(28422)	(27839)
White	0.797	0.843	0.772	0.784	0.790
	(0.402)	(0.364)	(0.420)	(0.412)	(0.408)
Black	0.138	0.114	0.156	0.143	0.138
	(0.345)	(0.318)	(0.363)	(0.350)	(0.345)
Other race	0.064	0.043	0.073	0.073	0.072
	(0.245)	(0.203)	(0.259)	(0.260)	(0.259)
Hispanic ethnicity	0.083	0.057	0.081	0.102	0.113
	(0.276)	(0.232)	(0.273)	(0.303)	(0.316)
Married at time of survey	0.728	0.794	0.697	0.701	0.712
	(0.445)	(0.404)	(0.460)	(0.458)	(0.453)
Observations	297,117	86,230	99,946	80,048	30,893

Note: The young-adult sample composition by survey year is as follows. The surveys that comprise the 19834-1991 sub-sample are: 78,759 observations from the 1990 census; 7,178 observations from the 2000 census; and 293 observations from the 2001 ACS. The surveys that comprise the 1992-1999 sub-sample are: 63,404 observations from the 2000 census; 5,976 observations from the 2001 ACS; 5,146 observations from the 2002 ACS; 4,933 observations from the 2003 ACS; 3,848 observations from the 2004 ACS; 6,975 observations from the 2005 ACS; 5,123 observations from the 2006 ACS; 2,997 observations from the 2007 ACS; 1,147 observations from the 2008 ACS; and 397 observations from the 2009 ACS. The surveys that comprise the 2000-2007 sub-sample are: 188 observations from the 2001 ACS; 548 observations from the 2002 ACS; 1,101 observations from the 2003 ACS; 1,825 observations from the 2004 ACS; 5,880 observations from the 2005 ACS; 7,670 observations from the 2006 ACS; 9,475 observations from the 2007 ACS; 9,309 observations from the 2008 ACS; 9,833 observations from the 2009 ACS; 8,970 observations from the 2010 ACS; 7,887 observations from the 2011 ACS; 6,262 observations from the 2012 ACS; 4,786 observations from the 2013 ACS; 3,321 observations from the 2014 ACS; 2,000 observations from the 2015 ACS; and 993 observations from the 2016 ACS. The surveys that comprise the 2008-2015 sub-sample are: 321 observations from the 2009 ACS; 993 observations from the 2010 ACS; 1,957 observations from the 2011 ACS; 3,124 observations from the 2012 ACS; 4,568 observations from the 2013 ACS; 5,707 observations from the 2014 ACS; 6,592 observations from the 2015 ACS; and 7,631 observations from the 2016 ACS.

Table 3. First-stage results: The effect of twins-first on the number of children by year of child's birth, 14 ≤ mother's age first child < 20.

	1976-2015	1976-1983	1984-1991	1992-1999	2000-2007	2008-2015
A. Overall sample						
Twins at first birth	0.808***	0.662***	0.842***	0.841***	0.855***	0.807***
	(0.024)	(0.043)	(0.042)	(0.038)	(0.043)	(0.059)
Partial R-squared	0.007	0.004	0.008	0.008	0.007	0.010
F-test	1074.72	237.585	400.099	487.542	392.149	182.814
Observations	313,921	64,779	67,634	102,968	14,729	3,283
B. White sample						
Twins at first birth	0.781***	0.585***	0.824***	0.832***	0.830***	0.877***
	(0.024)	(0.052)	(0.045)	(0.048)	(0.048)	(0.074)
Partial R-squared	0.0067	0.003	0.0086	0.0083	0.0067	0.0118
F-test	1039.73	125.957	325.641	298.150	296.163	138.176
Observations	211,977	47,171	45,617	67,120	41,492	9,656
C. Black sample						
Twins at first birth	0.856***	0.758***	0.856***	0.888***	0.957***	0.677***
	(0.037)	(0.076)	(0.077)	(0.051)	(0.093)	(0.102)
Partial R-squared	0.007	0.005	0.007	0.009	0.008	0.007
F-test	526.418	97.070	119.396	296.000	102.853	42.201
Observations	73,740	13,628	16,603	25,147	14,729	3,283

Note: All regressions include a quadratic for mother's age at first birth, a quadratic for the number of years since the first birth, survey year dummies, state of birth dummies, state/year unemployment rate, and a series of controls for race, ethnicity and marital status. Huber-White standard errors are shown from clustering by state of birth. * denotes statistical significant at 10%; ** denotes statistical significant at 5%; *** denotes statistical significant at 1%.

Table 4. First-stage results: The effects of twins-first on the number of children by year of child's birth, $20 \leq$ mother's age first child < 24.

	1984-2015	1984-1991	1992-1999	2000-2007	2008-2015
A. Overall sample					
Twins at first birth	0.746***	0.724***	0.738***	0.772***	0.760***
	(0.013)	(0.026)	(0.022)	(0.023)	(0.033)
Partial R-squared	0.012	0.012	0.011	0.011	0.013
F-test	3,438	755	1,086	1093	516
Observations	297,117	86,230	99,946	80,048	30,893
B. White sample					
Twins at first birth	0.724***	0.700***	0.708***	0.762***	0.761***
	(0.017)	(0.024)	(0.030)	(0.034)	(0.039)
Partial R-squared	0.011	0.012	0.010	0.010	0.012
F-test	1,868	814	561	497	375
Observations	236,931	72,672	77,112	62,750	24,397
C. Black sample					
Twins at first birth	0.825***	0.870***	0.778***	0.846***	0.815***
	(0.036)	(0.072)	(0.051)	(0.060)	(0.076)
Partial R-squared	0.016	0.016	0.014	0.016	0.019
F-test	511	142	228	195	112
Observations	41,108	9,827	15,581	11,437	4,263

Note: see table 3.

Table 5. IV results by year of child's birth, 14 ≤ mother's age first child < 20.

	1976-2015	1976-1983	1984-1991	1992-1999	2000-2007	2008-2015
A. Working for pay						
Overall	-0.069	-0.083***	-0.041*	-0.067***	-0.095***	-0.017
	(0.010)	(0.025)	(0.017)	(0.016)	(0.025)	(0.046)
White	-0.071***	-0.079*	-0.048*	-0.076***	-0.098**	-0.019
	(0.013)	(0.032)	(0.024)	(0.019)	(0.034)	(0.059)
Black	-0.069***	-0.070	-0.067*	-0.061*	-0.091*	0.030
	(0.019)	(0.037)	(0.030)	(0.027)	(0.036)	(0.109)
B. Number of weeks worked						
Overall	-2.651	-3.424***	-0.752	-3.195***	-3.207*	-0.164
	(0.507)	(0.990)	(0.835)	(0.755)	(1.262)	(2.305)
White	-2.404***	-2.198	-2.242	-3.050**	-2.374	1.394
	(0.589)	(1.500)	(1.171)	(0.944)	(1.516)	(2.718)
Black	-3.790***	-5.933***	-0.334	-4.120**	-5.008*	-3.138
	(0.982)	(1.801)	(1.656)	(1.527)	(2.325)	(5.022)
C. Number of hours worked						
Overall	-2.422***	-2.762**	-1.273*	-2.469***	-3.365***	-0.498
	(0.429)	(0.941)	(0.648)	(0.676)	(0.931)	(1.550)
White	-2.507***	-2.490	-1.419	-3.048***	-3.435**	0.219
	(0.497)	(1.307)	(0.979)	(0.780)	(1.147)	(2.453)
Black	-2.464***	-2.830*	-2.272*	-2.007	-3.136*	-2.113
	(0.719)	(1.300)	(1.119)	(1.131)	(1.280)	(3.760)
D. Mother's earned income						
Overall	-990.673***	-908.500	-376.642	-1428.750***	-866.048	-956.620
	(235.429)	(783.830)	(363.928)	(386.749)	(530.239)	(823.437)
White	-795.833**	-838.793	-307.057	-1141.761*	-751.759	-79.127
	(284.801)	(786.358)	(534.622)	(533.673)	(648.877)	(1064.450)
Black	-1437.555*	-1101.132	-933.826	-2271.112***	-593.037	-2466.267
	(558.385)	(2121.378)	(789.203)	(565.292)	(896.000)	(1811.084)
E. Family income						
Overall	-946.667*	-859.066	-619.553	-1523.700*	-741.654	428.214
	(388.483)	(1485.168)	(625.668)	(615.047)	(723.603)	(1865.798)
White	-1155.883	-1111.426	-770.134	-1845.200*	-1084.322	347.432
	(596.860)	(2529.655)	(850.724)	(812.982)	(964.530)	(2148.659)
Black	-809.492	-965.951	-591.889	-1962.886**	858.624	3982.840
	(652.746)	(2206.147)	(1385.974)	(706.037)	(1097.694)	(4626.961)

Note: see table 3.

Table 6. IV results by year of child's birth, 20 ≤ mother's age first child < 24.

	1984-2015	1984-1991	1992-1999	2000-2007	2008-2015
A. Working for pay					
Overall	-0.060***	-0.068**	-0.066***	-0.048***	-0.031
	(0.011)	(0.021)	(0.020)	(0.014)	(0.030)
White	-0.047**	-0.048	-0.053*	-0.039*	-0.040
	(0.016)	(0.027)	(0.026)	(0.016)	(0.032)
Black	-0.085***	-0.139***	-0.074*	-0.070	-0.008
	(0.017)	(0.031)	(0.033)	(0.043)	(0.047)
B. Number of weeks worked					
Overall	-3.525***	-4.495***	-3.985***	-2.486***	-1.051
	(0.456)	(0.887)	(0.947)	(0.731)	(1.527)
White	-3.157***	-3.935***	-3.385**	-2.415**	-2.025
	(0.625)	(1.141)	(1.254)	(0.859)	(1.635)
Black	-4.012***	-5.817***	-5.471***	-1.760	1.771
	(0.778)	(1.325)	(1.432)	(2.072)	(3.337)
C. Number of hours worked					
Overall	-2.689***	-3.347***	-2.769***	-2.261***	-1.143
	(0.390)	(0.799)	(0.745)	(0.558)	(1.008)
White	-2.033***	-2.462**	-2.147*	-2.036**	-0.588
	(0.529)	(0.946)	(0.947)	(0.662)	(1.331)
Black	-4.025***	-6.071***	-3.647**	-2.512	-2.947
	(0.884)	(1.334)	(1.365)	(1.957)	(2.199)
D. Mother's earned income					
Overall	-1776.154***	-2350.382***	-2055.522***	-1766.650***	996.151
	(235.630)	(594.577)	(517.344)	(409.417)	(1605.112)
White	-1636.810***	-2631.615***	-1564.279*	-1577.972**	637.408
	(286.320)	(511.073)	(701.430)	(517.092)	(1676.112)
Black	-1943.272**	-493.074	-3307.538***	-1599.908	-916.939
	(657.101)	(2286.638)	(815.272)	(989.736)	(1872.087)
E. Family income					
Overall	-1830.685***	-2141.846*	-2432.250**	-1996.010**	2082.437
	(427.406)	(974.159)	(789.374)	(751.789)	(2441.737)
White	-1875.130***	-3022.690**	-1642.426	-2395.425**	2433.251
	(516.033)	(938.666)	(1166.033)	(859.992)	(2903.081)
Black	-1631.596	1040.145	-4191.784***	-800.695	-1908.603
	(848.047)	(2658.871)	(1118.835)	(1367.754)	(1828.396)

Note: see table 3.

Appendix

Table 1A. Variable means for teenage mothers by employment status.

	Overall sample Twin mothers	Working for pay Twin mothers	Overall sample Non-twin mothers	Working for pay Non-twin mothers
Number of children	2.979	2.857	2.210	2.107
	(1.048)	(0.976)	(1.023)	(0.973)
Age at 1st birth (years)	18.024	18.028	17.898	17.902
	(1.323)	(1.330)	(1.314)	(1.315)
Age at survey (years)	25.345	25.546	25.489	25.581
	(3.105)	(3.044)	(3.060)	(3.052)
Years since 1st birth	7.498	7.698	7.812	7.906
	(3.431)	(3.411)	(3.356)	(3.355)
Worked for pay	0.681	1.000	0.734	1.000
	(0.466)	(0.000)	(0.442)	(0.000)
No. of weeks worked	26.374	38.721	28.630	39.005
	(22.463)	(16.206)	(22.158)	(16.255)
No. of hours worked	24.343	35.740	26.136	35.608
	(18.737)	(10.393)	(18.294)	(10.894)
Mom's income (1999 dollars)	8938	13123	9758	13294
	(11891)	(12356)	(12371)	(12708)
Family income (1999 dollars)	26071	28871	27545	29542
	(22848)	(23627)	(24387)	(24519)
White	0.621	0.612	0.676	0.667
	(0.485)	(0.487)	(0.468)	(0.471)
Black	0.285	0.295	0.234	0.246
	(0.451)	(0.456)	(0.424)	(0.431)
Other race	0.094	0.093	0.090	0.088
	(0.292)	(0.291)	(0.286)	(0.283)
Hispanic ethnicity	0.114	0.111	0.122	0.118
	(0.318)	(0.314)	(0.327)	(0.323)
Married at time of survey	0.576	0.543	0.599	0.570
	(0.494)	(0.498)	(0.490)	(0.495)
Observations	2,882	1,963	311,039	228,303

Table 2A. Variable means for young-adult mothers by employment status.

	Overall sample Twin mothers	Working for pay Twin mothers	Overall sample Non-twin mothers	Working for pay Non-twin mothers
Number of children	2.586	2.509	1.865	1.767
	(0.781)	(0.733)	(0.818)	(0.767)
Age at 1st birth (years)	21.775	21.770	21.704	21.699
	(1.127)	(1.131)	(1.137)	(1.141)
Age at survey (years)	26.735	26.831	26.758	26.789
	(1.902)	(1.897)	(1.918)	(1.927)
Years since 1st birth	5.199	5.298	5.313	5.348
	(2.198)	(2.194)	(2.220)	(2.240)
Worked for pay	0.705	1.000	0.741	1.000
	(0.456)	(0.000)	(0.438)	(0.000)
No. of weeks worked	28.189	39.989	30.456	41.093
	(22.632)	(15.957)	(22.233)	(15.159)
No. of hours worked	24.481	34.729	26.056	35.157
	(18.457)	(11.281)	(18.271)	(11.421)
Mom's income (1999	10427	14792	11570	15611
dollars)	(14216)	(14904)	(13505)	(13528)
Family income (1999	33656	35502	36095	37496
dollars)	(26821)	(26563)	(28021)	(27097)
White	0.751	0.742	0.798	0.783
	(0.433)	(0.438)	(0.402)	(0.412)
Black	0.186	0.199	0.138	0.153
	(0.390)	(0.399)	(0.345)	(0.360)
Other race	0.063	0.060	0.064	0.064
	(0.243)	(0.237)	(0.245)	(0.245)
Hispanic ethnicity	0.073	0.069	0.083	0.083
	(0.260)	(0.254)	(0.276)	(0.276)
Married at time of	0.688	0.653	0.728	0.688
survey	(0.463)	(0.476)	(0.445)	(0.463)
Observations	3,443	2,427	293,674	217,650

Table 3A. OLS regressions by year of child's birth, 14≤ mother's age first child < 20.

	1976-2015	1976-1983	1984-1991	1992-1999	2000-2007	2008-2015
A. Working for pay						
Overall	-0.093***	-0.110***	-0.092***	-0.083***	-0.082***	-0.086***
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.006)
White	-0.104***	-0.111***	-0.112***	-0.096***	-0.097***	-0.100***
	(0.002)	(0.003)	(0.003)	(0.002)	(0.003)	(0.008)
Black	-0.070***	-0.103***	-0.064***	-0.054***	-0.051***	-0.063***
	(0.002)	(0.004)	(0.003)	(0.003)	(0.004)	(0.008)
B. Number of weeks worked						
Overall	-5.636***	-6.117***	-5.582***	-5.380***	-5.125***	-5.843***
	(0.064)	(0.121)	(0.097)	(0.078)	(0.093)	(0.218)
White	-6.220***	-6.370***	-6.620***	-6.061***	-5.867***	-6.429***
	(0.082)	(0.127)	(0.128)	(0.121)	(0.135)	(0.349)
Black	-4.339***	-5.373***	-4.082***	-3.901***	-3.662***	-4.772***
	(0.065)	(0.139)	(0.129)	(0.107)	(0.149)	(0.468)
C. Number of hours worked						
Overall	-3.901***	-4.584***	-3.830***	-3.594***	-3.422***	-3.576***
	(0.062)	(0.102)	(0.091)	(0.075)	(0.096)	(0.205)
White	-4.332***	-4.653***	-4.596***	-4.154***	-4.028***	-4.063***
	(0.068)	(0.119)	(0.113)	(0.085)	(0.104)	(0.301)
Black	-2.920***	-4.142***	-2.718***	-2.405***	-2.226***	-2.718***
	(0.058)	(0.124)	(0.119)	(0.104)	(0.170)	(0.284)
D. Mother's earned income						
Overall	-2633.650***	-2684.324***	-2555.301***	-2859.130***	-2389.008***	-2145.936***
	(58.945)	(93.995)	(84.686)	(65.762)	(70.637)	(126.394)
White	-2870.609***	-2805.159***	-2886.106***	-3134.143***	-2704.851***	-2312.864***
	(52.285)	(94.311)	(75.167)	(69.152)	(81.090)	(154.773)
Black	-2075.325***	-2286.427***	-1988.132***	-2205.944***	-1720.524***	-1957.849***
	(60.517)	(94.601)	(124.639)	(82.391)	(100.437)	(205.019)
E. Family income						
Overall	-2577.563***	-2466.078***	-2380.342***	-2997.038***	-2330.053***	-1677.075***
	(79.738)	(104.764)	(108.812)	(108.363)	(123.951)	(219.368)
White	-2869.772***	-2767.907***	-2756.797***	-3341.936***	-2676.010***	-1514.985***
	(98.863)	(118.215)	(130.491)	(155.199)	(141.590)	(305.164)
Black	-1865.805***	-1866.581***	-1740.756***	-2099.232***	-1587.319***	-1742.842***
	(47.250)	(125.088)	(150.766)	(101.778)	(138.361)	(252.792)

Note: see table 3.

Table 4A. OLS regressions by year of child's birth, 20 ≤ mother's age first child < 24.

	1984-2015	1984-1991	1992-1999	2000-2007	2008-2015
A. Working for pay					
Overall	-0.114***	-0.128***	-0.107***	-0.111***	-0.108***
	(0.002)	(0.003)	(0.002)	(0.003)	(0.004)
White	-0.121***	-0.130***	-0.116***	-0.119***	-0.117***
	(0.002)	(0.003)	(0.003)	(0.004)	(0.005)
Black	-0.083***	-0.110***	-0.069***	-0.074***	-0.063***
	(0.003)	(0.005)	(0.004)	(0.005)	(0.009)
B. Number of weeks worked					
Overall	-6.801***	-7.467***	-6.575***	-6.451***	-6.494***
	(0.089)	(0.110)	(0.095)	(0.157)	(0.250)
White	-7.080***	-7.662***	-6.904***	-6.847***	-6.826***
	(0.110)	(0.127)	(0.106)	(0.191)	(0.289)
Black	-5.514***	-6.440***	-5.266***	-4.799***	-4.865***
	(0.130)	(0.253)	(0.201)	(0.294)	(0.348)
C. Number of hours worked					
Overall	-5.081***	-5.554***	-4.903***	-4.904***	-4.692***
	(0.058)	(0.086)	(0.082)	(0.095)	(0.152)
White	-5.363***	-5.659***	-5.311***	-5.256***	-5.084***
	(0.056)	(0.094)	(0.086)	(0.101)	(0.175)
Black	-3.686***	-4.783***	-3.197***	-3.337***	-2.662***
	(0.099)	(0.195)	(0.165)	(0.233)	(0.323)
D. Mother's earned income					
Overall	-3355.889***	-3396.761***	-3512.957***	-3309.011***	-2781.149***
	(50.522)	(92.919)	(74.802)	(58.821)	(93.270)
White	-3414.223***	-3408.424***	-3604.171***	-3380.604***	-2899.827***
	(47.940)	(83.114)	(75.744)	(77.045)	(117.206)
Black	-2939.451***	-3098.265***	-2937.368***	-2946.077***	-2146.963***
	(140.475)	(190.232)	(193.849)	(186.754)	(226.793)
E. Family income					
Overall	-2193.535***	-2051.462***	-2454.997***	-2134.056***	-1857.520***
	(111.530)	(116.790)	(103.586)	(188.936)	(223.711)
White	-2106.625***	-1935.303***	-2424.298***	-2023.105***	-1831.713***
	(115.789)	(127.578)	(113.612)	(211.600)	(270.782)
Black	-2591.649***	-2633.764***	-2591.005***	-2646.097***	-2015.888***
	(131.196)	(195.701)	(190.553)	(240.286)	(363.101)

Note: see table 3.