INCOME INEQUALITY AND EARLY NON-MARITAL CHILDBEARING:
AN ECONOMIC EXPLORATION OF THE "CULTURE OF DESPAIR"

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ABSTRACT

Using individual-level data from the United States and a number of other developed countries, we empirically investigate the role of income inequality in determining rates of early, non-marital childbearing among low socioeconomic status (SES) women. We present robust evidence that low SES women are more likely to give birth at a young age and outside of marriage when they live in more unequal places, all else held constant. Our results suggest that inequality itself, as opposed to other correlated geographic factors, drives this relationship. We calculate that differences in the level of inequality are able to explain a sizeable share of the geographic variation in teen fertility rates both across U.S. states and across developed countries. We propose a model of economic "despair" that facilitates the interpretation of our results. It reinterprets the sociological and ethnographic literature that emphasizes the role of economic marginalization and hopelessness into a parsimonious framework that captures the concept of "despair" with an individual's perception of economic success. Our empirical results are consistent with the idea that income inequality heightens a sense of economic despair among those at the bottom of the distribution.

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I. INTRODUCTION

Rates of early, non-marital childbearing vary tremendously across countries and across states. The United States is consistently at the high end of this distribution, with a rate of teen childbearing that greatly exceeds that of other developed countries. Within the United States the experience is far from uniform. Some states have teen childbearing rates that are roughly comparable to those found in Europe, whereas others have rates that are over three times that level.

These patterns are documented in Figures 1 and 2, which display teen childbearing rates by country and by states within the U.S., respectively. The teen birth rate of 41.5 per 1,000 in the United States is a multiple of the level that exists in other developed countries. For example, the rate is 25.9 in the United Kingdom, 14.1 in Canada, and 4.3 in Switzerland. Figure 2 shows that tremendous variation exists across states as well: some states have rates that are comparable to those in other developed countries, but others have extremely high rates. For example, the rates in Texas, New Mexico, and Mississippi (over 60 per 1,000) are more than three times the rates in New Hampshire, Massachusetts, and Vermont.[1] As we document subsequently, these geographic differences are long-standing and persistent. We consider this geographic variation to pose both a challenge and potentially a clue in thinking about factors that may be important to understanding rates of early non-marital childbearing.

Despite over four decades of research on the topic of early, non-marital childbearing, the best available research to date has very little to say about why these differences exist and why they are so persistent. Why is it that teenagers in the United States are so much more likely than

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[1] This cross-state variation in teen birth rates does not simply reflect cross-state variation in overall birth rates. The highest fertility rates to women age 15-44 are found in Alaska, Idaho, and Utah, which rank 35, 28, and 19, respectively, in terms of teen birth rates (Martin et al., 2010). In addition, this cross-section variation is substantially larger than recent time-series variation, which has garnered so much attention. Between 1991 and 2008, the national teen birth rate fell from a peak of 62 to 42.

their counterparts in other countries to give birth when young and unmarried? Why are teens in some parts of the United States so much more likely to have a teen birth than their counterparts in other parts of the country? In this paper we argue that the standard explanations studied by economists come up short of being able to explain any substantial share of the variation. We propose a new direction of research and present robust empirical evidence that income inequality can explain a substantial share of this geographic variation.

A striking correlation exists between income inequality and aggregate rates of teen childbearing, both across countries and across states. But places with more inequality also differ in many other dimensions that could affect rates of teen childbearing. To determine whether aggregate inequality measures have a causal relationship with rates of early non-marital childbearing, we conduct an empirical analysis of individual level data that examines whether income inequality operates on individuals likely to be affected in a negative way – namely, those at the bottom of the income distribution. We use individual level data from the National Survey of Family Growth (NSFG) to look across states in the U.S. and individual level data from the Fertility and Family Survey (FFS), conducted by the United Nations, to look across several developed countries. We find that women who grew up in low socioeconomic circumstances have more teen, non-marital births when they live in higher inequality locations, all else equal. The proximate mechanism driving this finding is less frequent use of abortion.

Of course, other geographic characteristics are likely to be correlated with inequality and may limit our ability to draw causal conclusions from this relationship. Our analysis controls for demographic characteristics of the population and a broad array of public policies that could otherwise pose problems for our interpretation. We also show that our results are unchanged when we hold constant other economic factors, like the absolute income levels of the poor, as

well as other potentially important environmental factors such as social capital or political climate. While we could never completely rule out the existence of an omitted factor, the robustness of the relationship is striking.

We propose a model of economic "despair" to rationalize these results. Our model is heavily influenced by the existing ethnographic and sociological literature on the topic, including the work of Clark (1965), Lewis (1969), Wilson (1987), and Edin and Kefalas (2005), among others. Despite specific differences, all these authors broadly view despair and hopelessness as playing a key role in driving rates of non-marital childbearing among the poor. Our contribution to this literature is to capture this qualitative idea in a parsimonious model within the economics paradigm of a utility-maximizing framework. We explicitly treat an early non-marital birth as being the result of individual decision-making.

We focus on the idea that it is the perceived inability of poor women to improve their situation through work or marriage that leads them to choose motherhood when young and single, rather than delaying until a later period. When a poor young woman perceives that socioeconomic success is unachievable to her, she is more likely to embrace motherhood in her current position, as there is little option value to be gained by delaying the immediate gratification of having a baby. When there is relatively more hope of economic advancement – a perception that is more likely in a more equal or mobile society – it is relatively more desirable to delay motherhood and invest in human or social capital.[2] The operational link to the empirical work is that the combination of being poor and living in a more unequal (and less mobile)

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[2] We show subsequently that inequality and mobility are highly correlated across societies. We also discuss that inequality theoretically could have an opposing "aspirational" effect (more to strive for in unequal societies), so that the ultimate impact of inequality on teen childbearing outcomes is an empirical question. The evidence is consistent with our interpretation here.

society contributes to the decision to have a baby when young and unmarried. Our empirical results are consistent with these ideas.

Our scholarly contribution is twofold. First, we present empirical evidence supporting the role of income inequality in driving rates of early, non-marital childbearing among those at the bottom of the income distribution. Our estimates suggest that inequality can explain a sizable share of the geographic variation observed in teen childbearing rates, on the order of 10 to 50 percent. To date, no other explanation can come close to explaining as much of the geographic variation.

Second, we believe our economic model of despair constitutes an important contribution to the literature in economics that considers so-called risky behaviors. In our conceptualization, these behaviors might be more appropriately considered "drop out" behaviors, in so far as adolescents are choosing to drop out of the mainstream climb to socio-economic success. Though this paper is focused entirely on early non-marital childbearing, we believe our model is applicable to a number of other contexts that involve current benefits and future economic costs.

II. REVIEW OF PREVIOUS RESEARCH

The standard economics model of childbearing considers an individual who maximizes utility over children and other consumption subject to a budget constraint (cf. Becker and Lewis, 1973). Preferences are generally assumed to be given and stable, and explanations have focused on differences in constraints, often generated by particular policies and institutions. Economists have tended to explore the relevance of factors such as the incentives of the welfare system, the role that abortion policy plays, the impact of labor market conditions, and the like. Other research has examined the impact of an individual's own economic disadvantage and her likelihood of having an early non-marital birth.

The political scientist Charles Murray (1986) wrote in his now-famous book, *Losing Ground*, that the welfare system provided incentives for couples to have a child outside of marriage by reducing both the financial rewards of marriage and the financial costs of out-of-wedlock childbearing. This hypothesis became politically popular among conservatives and helped usher in an era of welfare reform. It also spawned a vast empirical literature in economics investigating the issue. Moffitt (1998 and 2003) provides an overview of the large literature on the topic, concluding that more generous welfare benefits likely have only a modest positive effect on rates of non-marital childbearing. With regard to the variation across countries, the lower rate of teen childbearing in Europe with its much more generous welfare system provides a *prima facie* case against the hypothesis that social support is largely to blame for high rates of teen childbearing in the United States.

Economists have also examined a host of other policy and institutional factors relevant to the costs of avoiding or not avoiding a non-marital or teen birth. A highly incomplete list of such studies includes previous work that we have conducted elsewhere on the effect of various policies and environmental conditions, such as restrictive abortion policies (Levine, Trainor, and Zimmerman, 1996; Levine, 2003); welfare reform (Kearney, 2004); labor market conditions (Levine, 2001) and access to affordable contraception (Kearney and Levine, 2009a). These empirical studies have generally found that changes in such "prices" do have impacts on teen and non-marital childbearing, but individually these factors can account for only very small shares of the total variation in non-marital childbearing.

Moving away from traditional economic models of fertility, Akerlof, Yellen, and Katz (1996) propose a "technology shock" hypothesis for the rise in non-marital childbearing in the U.S. in the later 20th century. They relate the erosion of the custom of "shotgun marriage" – the

practice of getting married between conception and birth -- to the legalization of abortion and the increased availability of contraception to unmarried women in the United States. The story is one of decreased bargaining power on the part of women who do not adopt either birth control or abortion. This theory is an intriguing explanation for the decrease in shotgun marriages in the U.S. over the relevant decades, but it is unlikely to have much explanatory power for the geographic variation in outcomes since those technology shocks happened everywhere.

Behavioral economists O'Donohue and Rabin (1999) suggest that teens are "hyperbolic discounters," who place disproportionate weight on present happiness as compared to future well-being. Other scholars suggest that teen childbearing is attributable to teens' stage of cognitive development, arguing that they are not quite ready to make the types of decisions that would prevent a pregnancy (for example, Brooks-Gunn and Furstenberg, 1989; Hardy and Zabin, 1991; Brooks-Gunn and Paikoff, 1997).While limited decision-making capacity surely is an issue for some set of teens, we note that these claims have an element of universality to them that cannot begin to explain the striking differences in rates of early non-marital childbearing across socioeconomic groups, over time, or across states or countries. In other words, we doubt that the particularly high rate of teen childbearing among U.S. teens as compared to their counterparts in Europe can be attributed to the more limited decision making capacity -- or more present-biased preferences -- of the teenage brain in America.

Another line of research considers the relationship between background disadvantage and rates of early childbearing (cf. Duncan and Hoffman, 1990; An, Haveman, and Wolfe, 1993; Lundberg and Plotnick, 1995; and Duncan, et al., 1998). It is well-known that growing up in disadvantaged circumstances, such as in poverty or to a single mother, is associated with much higher rates of early childbearing. In a previous examination of cohort rates of early

childbearing, we find that the proportion of a female cohort born economically disadvantaged – as captured by being born to a teen mother, a single mother, or to a mother with a low level of education – is tightly linked to the subsequent rate of early childbearing in that cohort (Kearney and Levine, 2009b). But, strikingly, we find that state and year of birth fixed effects capture much of the variation. We interpret that finding as suggestive of the importance of some "cultural" dimension, otherwise un-modeled in that framework; we return to such an alternative perspective subsequently.

Although economists have been contributing to discussions of early, non-marital childbearing for several decades, the first contributors to the discussion were other social scientists. Their work pursued a parallel, rarely (if ever) intersecting track. Daniel Patrick Moynihan's 1965 report first drew attention to the issue of non-marital childbearing among black families in the U.S., when the rate was one in three.[3] At about the same time, the social theories of the psychologist Clark (1965) and the anthropologist Lewis (1969) - who developed the theory of the "Culture of Poverty" - were met with controversy. Lewis (1969) wrote the following:

> *The culture of poverty is both an adaptation and a reaction of the poor to their marginal position in a class-stratified, highly individuated, capitalistic society. It represents an effort to cope with the feelings of hopelessness and despair that develop from the realization of the improbability of achieving success in terms of the values and goals of the larger society ... [M]any will tell you that marriage by* law, by the church, or by both is the idea form of marriage; but few marry ... Women often turn down offers of marriage because they feel that it ties them down to men who are immature, punishing, and generally unreliable (p. 189-190)

This lack of opportunity was observed explicitly by Clark (1965), who wrote the following:

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[3] Moynihan (1965) argued that the deterioration of the nuclear family, and the rise of the female-headed households, was hindering the economic progress of blacks in the U.S. As a policy matter, he argued that it was crucial to improve the job prospects of black men in order to keep them engaged in the family and community as father figures, and thereby curb the steady increase in rates of out-of-wedlock childbearing and divorce that was contributing to increased rates of poverty among black communities.

> *In the ghetto, the meaning of the illegitimate child is not ultimate disgrace. There is not the demand for abortion or for surrender of the child that one finds in more privileged communities. In the middle class, the disgrace of illegitimacy is tied to personal and family aspirations. In lower-class families, on the other hand, the girl loses only some of her already limited options by having an illegitimate child; she is not going to make a "better marriage" or improve her economic and social status either way. On the contrary, a child is a symbol of the fact that she is a woman, and she may gain from having something of her own. Nor is the boy who fathers an illegitimate child going to lose, for where is he going? The path to any higher status seems closed to him in any case (p. 72).*

The sociologist Wilson (1987) revived serious scholarship on the topic with his book *The Truly Disadvantaged.* The distinction between Wilson and Clark is largely the focus on the lack of jobs itself in Wilson, not the social attitude that results from the lack of jobs. Either way, the lack of opportunity is what is driving the childbearing outcomes in both viewpoints, as described in the following excerpt:

> *Thus, in a neighborhood with a paucity of regularly employed families and with the overwhelming majority of families having spells of long-term joblessness, people experience a social isolation that excludes them from the job network system that permeates other neighborhoods and that is so important in learning about or being recommended for jobs that become available in various parts of the city ... Moreover, unlike the situation in earlier years, girls who become pregnant out of wedlock invariably give birth out of wedlock because of a shrinking pool of marriageable, that is, employed black men (p. 57).*

The relevant environmental factor for women in this argument is the weak marriage market that is attributable to the lack of jobs for men in the inner city. Wilson is clear to point out that his focus is on "social isolation" and not the "culture of poverty."

More recently, Edin and Kefalas (2005) contributed an influential ethnographic account of non-marital childbearing among poor women. They make the following observation:

> *... the extreme loneliness, the struggles with parents and peers, the wild behavior, the depression and despair, the school failure, the drugs, and the general sense that life has spun completely out of control. Into this void comes a pregnancy and then a baby, bringing the purpose, the validation, the companionship, and the order that young women feel have been so sorely lacking. In some profound*

sense, these young women believe, a baby has the power to solve everything (p. 10).

From this perspective, having a baby at a young age outside the scope of marriage is not the result of a constraint, but something that the woman values.

Our reading of these seminal and influential works is that they find common ground in the notion that growing up in an environment where there is little chance of social and economic advancement leads young women to bear children outside of marriage. These women perceive that they have so little chance for success in life not solely because of their own disadvantage, but also because of the environment in which they live. They see no reason to postpone having a child and may even benefit from having one, regardless of marital status.[4]

This literature provides potentially useful insights regarding why some places have so much higher rates of early non-marital childbearing than others. Disadvantaged individuals who live in locations with little opportunity for economic advancement will be more likely to have an early, non-marital birth. We will return to these ideas subsequently in interpreting the results of our econometric analysis.

III. ECONOMETRIC APPROACH

A potential clue in understanding the determinants of early, non-marital childbearing is the persistence in its geographic patterns. Year-to-year and even decade-to-decade variability within a location is very limited. The correlation in teen birth rates within states, for instance, between 1980 and 2008 is 0.92. This suggests that longstanding differences across states are critical. In terms of thinking about a typical analysis of state-year level regressions, statistically

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[4] Many of these arguments, and particularly the earlier ones, focus directly on the issues of race. The Moynihan Report (1965) which first broadly publicized the issue of rising non-marital fertility, also focused on race. At the time, one in three births to black women was outside of marriage whereas the rate for whites was much lower. Now that rates of early, non-marital childbearing are high for all women (albeit still higher for black women), our view is that this is less an issue of race today than it used to be.

the state fixed effect is paramount in explaining variation in teen childbearing research. In this research we are trying to see whether we can get inside the black box of that fixed effect.

One potential culprit among many is income inequality, which tends to be fairly persistent within a place and shows a strong correlation with rates of early, non-marital childbearing.[5] Wilkinson and Pickett (2009) call attention to the correlation between inequality and a broad range of "social ills," including increased crime, drug use, mortality, and teen childbearing, among others. Figures 3 and 4 present scatter plots displaying these relationships across developed countries and across states in a similar manner to their analysis. The correlation is positive, and is especially large in the case of cross-country comparisons. This correlation is noteworthy, but it does not necessarily imply a causal positive relationship between income inequality and early, non-marital childbearing. [6]

An important step in establishing a causal relationship is to determine that inequality operates on those who are most likely to be negatively impacted by it. With this purpose in mind, our primary empirical test is based on determining whether low SES women in high inequality locations are more likely to have an early, non-marital birth compared to higher SES women in those locations. We estimate regression models for a series of fertility outcomes (birth, conception, etc.) controlling for year-of-birth fixed effects and state fixed effects. We also control for individual level demographics, public policy variables and time-varying labor market conditions in these models.

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[5] Using data from the 1980 and 2000 Censuses, which we describe subsequently, we find that the correlation in the 50/10 ratio across states between those years is 0.74.

[6] Wilkinson and Pickett (2009) focuses on correlational relationships, which are merely the starting point of our empirical investigation. Furthermore, the robustness of their correlational relationships has been called into question in other contexts, for example, see Deaton and Lubotsky (2003) on the relationship between inequality and mortality.

The key variable in our models is the interaction between long-term measures of inequality and a woman's socioeconomic status. We are interested in explaining persistent differences in rates of early, non-marital childbearing across places. We therefore focus our analysis on "fixed" characteristics of environments. In terms of inequality, this means we empirically consider variation in long-term averages in state and national measures of income inequality. Note that these long-term inequality measures are perfectly correlated with the state fixed effects, and we thus do not separately estimate the conditional main effect.

More formally, we estimate regression models for multiple outcomes by age 20 of the form:

$$Outcome_{isc} = \beta_0 + \beta_1\left(I_s \cdot LS_{isc}\right) + \beta_2 LS_{isc} + \beta_3 X_{isc} + \beta_4 E_{sc} + \gamma_s + \gamma_c + \varepsilon_{isc} \qquad [1]$$

where I is our measure of inequality, LS is an indicator of low socioeconomic status, and the interaction term is the regressor of primary interest. The subscript i indexes individuals, s indexes states (or countries), and c indexes birth cohorts. The terms γ_s and γ_c represent state (or country) and birth cohort fixed effects, respectively. The vector *X* consists of additional personal demographic characteristics – age, age squared, race/ethnicity, and an indicator for living with a single parent at age 14.

The vector *E* captures environmental factors including relevant public policies and labor market conditions in the state-year: the unemployment rate, an indicator for a welfare family cap, the maximum welfare benefit for a family of three, an indicator for SCHIP implementation, an indicator for whether the state Medicaid program covers abortion, an indicator for whether state abortion regulations include parental notification or mandatory delay periods, and whether the state Medicaid program includes expansion policies for family planning services (see Kearney and Levine, 2009a for a discussion of these policies. The data sources used to create these

variables are described in the data appendix. By including all of these individual and state level controls in the model, our estimated effect of inequality for low-SES women is net of effects driven by policies that might be correlated with inequality.

Our primary question of interest is whether β_1 is positive: are low-SES women in high inequality states relatively more likely to have a non-marital birth by age 20? We consider the multiple channels through which a difference in birth rates could be realized. First, an individual can take actions with regard to sexual behavior and contraceptive practices; low SES women in more unequal places may be more likely to get pregnant. Second, a non-marital childbirth could be avoided through the choice to end a pregnancy; low SES women in more unequal places may be relatively less likely to choose an abortion to end her pregnancy. And finally, non-marital births depend upon the parents choosing to remain unmarried after a pregnancy occurs; low SES women in more unequal places may be relatively less likely to get married before the birth through a so-called "shot-gun" marriage.

One limitation of this modeling approach is that any other characteristic of the socioeconomic environment that might be correlated with inequality and also directly related to non-marital early childbearing propensities among low-SES women will be captured by our long-run inequality/SES interaction term. Although it is impossible to completely rule out this form of omitted variable bias, we examine whether introducing additional interactions of SES with other potentially troublesome characteristics change our results. More formally, we estimate "horse race" models that take the form:

$$Outcome_{isc} = \beta_0 + \beta_1\left(I_s \cdot LS_{isc}\right) + \beta_2 LS_{isc} + \beta_4\left(A_s \cdot LS_{isc}\right) + \beta_4 X_{isc} + \beta_5 E_{sc} + \gamma_s + \gamma_c + \varepsilon_{isc} \qquad [2]$$

This specification is largely the same as that in equation [1] except that we now include alternative factors (A) that are time invariant within states (not countries, as we describe subsequently) and we interact them with an indicator for low SES.

We consider two categories of alternatives. The first includes other features of the income and wage distribution. The second includes other social measures that are more directly focused on identifying alternative explanations – political composition of the state, a measure of religiosity of the state, the percentage of the population that is minority, the incarceration rate, poverty rate, and Putnam's Social Capital Index. We describe all of the alternatives we consider in more detail when we report the results of our analysis.

IV. DATA DESCRIPTION

Our empirical analysis takes advantage of cross-state variation in outcomes using five waves of data from the National Survey of Family Growth (NSFG) and cross-national variation in outcomes using data from the Fertility and Family Surveys (FFS).

A. NSFG Data

The NSFG has evolved considerably since its inception in 1973. Initially, it focused exclusively on married women between the ages of 15 and 44. Beginning with the 1982 survey, all women in this age range were included regardless of marital status; we restrict our attention to the surveys since then. The 2002 survey was also the first to include men in this age range, but with just two waves of men available and a relatively small sample size for state level analyses, these data are insufficient to study their childbearing outcomes for this project. Also, beginning in 2006, the survey changed its design from one that was conducted every six or seven years to one that is conducted annually, but with smaller samples in each year. In the end, we use data

from the 1982, 1988, 1995, 2002, and 2006-2008 surveys. These surveys provide observations for over 42,000 women between the ages of 15 and 44.

In all of our subsequent analysis, we restrict our attention to women who have turned age 20 (or 25 where appropriate) in 1976 or afterwards. The age restriction follows naturally from our outcome measures that are measured at age 20 or 25.   We impose the year restriction to avoid much of the social and behavioral changes that were associated with the introduction and diffusion of the birth control pill in the 1960s and abortion legalization in the early 1970s. After imposing these sample restrictions, we still have nearly 27,000 observations in these data.

Each survey contains complete pregnancy histories, which we can use to generate measures of pregnancies and pregnancy resolution (including childbearing) by age 20 (that is, through age 19). One potential problem with these data is the reporting of abortions. A woman who has had an abortion may report it accurately, report it as a miscarriage, or not report the pregnancy at all. For our purposes, we focus on the incidence of "pregnancy failure," which includes either a miscarriage or an abortion. This measure should capture behavioral changes in abortion regardless of how they are reported as long as the pregnancy itself is reported. Although miscarriages may not be purely biologically determined (Ashcraft and Lang, 2010), we believe it is reasonable to assume that the vast majority of movements in pregnancy failures are generated from changes in abortion decisions (particularly since fetal deaths are so rare).

Data on age at first marriage is also available so that we can ascertain whether the observed pregnancies occurred before marriage.   We can also approximate whether the pregnancy led to a marriage that occurred before the birth of the child. We define these so-called "shotgun marriages" as a birth that follows a marriage by six months or less.

The importance of using microdata for this exercise is that we are able to link fertility histories to personal characteristics. In particular, the hypothesis we test is about the impact of inequality on the fertility decisions of those with low socioeconomic status. It is critical to be able to provide an operational definition of low socioeconomic status and identify women who satisfy it. Although ideally we would have access to family income when women were growing up, the retrospective nature of the data prohibits that. Instead, we categorize women according to their mother's level of education, focusing on the children of high school dropouts as the ones who should be affected the most when they grow up in locations with greater inequality. Another important feature of these data is the availability of state identifiers, which enable us to make the link to the level of income inequality where the women grew up.[7]

We attach to these outcome measures a set of environmental factors that existed in the respondent's state of residence and the year in which she turned age 19; when we analyze birth outcomes by age 25, we attach state-level factors for the year in which a woman turned age 24. These factors include the policy variables listed above -- the level of welfare generosity and the status of welfare reform, SCHIP implementation, Medicaid family planning expansions, and the types of abortion restrictions. In addition, we attach the state-level unemployment rate in that year. We additionally attach alternative characteristics measures that we use in our horse race specifications – percentage of votes to Democrats, Index of religiosity, percentage of population that is minority, incarceration rate, poverty rate, and Putnam's Social Capital Index.

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[7] State identifiers are available in the NSFG for researchers with special permission from the National Center for Health Research. We accessed these data on-site at the NCHS Research Data Center in Hyattsville, MD. These state identifiers focus on the respondent's state of residence at the time of the survey, which may be different from their state of residence during their teenage years. We have to assume that any interstate migration related to the interaction of inequality and socioeconomic status during the intervening years is small.

Of primary interest, we attach measures of income inequality in the respondent's state of residence.[8] We created these measures using microdata from the 1980, 1990, and 2000 Censuses along with the 2006-2008 American Community Surveys on household income. These data are available from IPUMS-USA (Ruggles, 2010). Using these data, we estimated income cut-offs at the 10th and 50th percentiles for each state and survey year and then generate ratios of these measures (i.e. 50/10 ratio) as our indicator of inequality. We use these data to estimate long-run averages in measured inequality within states.

Table 1 tabulates inequality by state, which varies widely. Interestingly, the fifth highest inequality state by this measure is Massachusetts with a 50/10 ratio of 4.6 and the fifth lowest inequality state is its neighbor, Vermont, with a 50/10 ratio of 3.6. This suggests that as a rough gauge, moving from a low inequality state to a high inequality state increases the 50/10 ratio by around 1. We will use this number in interpreting our results.

Figures 5 and 6 present some descriptive statistics from the NSFG data that provide some useful background for interpreting our subsequent analysis. In Figure 5, we present trends in rates of childbearing by age 20 for all women and by marital status at the time of first birth. This figure shows that the rate of early childbearing has fluctuated around a level of about 20 percent since the mid-1970s. The most recent trend is downward; of those who turned age 20 in 2006 around 17 percent had already given birth. This relative stability masks dramatic differences in

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[8] We have considered whether the state is the right level of aggregation for this analysis. On the one hand, teens may have a better perspective on the economic well-being of those more immediately around her. On the other, the broader environment may better capture for her what her available opportunities are. Data issues also hinder our ability to conduct an analysis at local levels. First, our geographic identifier in the NSFG is the location of current residence and mobility is a much greater concern at the local level than at a broader level. Second, the only sub-state identifier available even in restricted NSFG data is the county of residence. This limits the sample size available for our analysis because public use census data only identifies county of residence for 40 percent of the population. Nevertheless, we have conducted a county-level analysis identical in form to that reported here using the subset of NSFG respondents who live in counties for which Census income data are available. The results are qualitatively similar than those reported here, but smaller in size. Rather than a coefficient of 0.053 (0.015) in Column 1 of Table 3, the comparable coefficient with county data is 0.031 (0.012). This is consistent with the attenuation bias that would result from cross-county migration.

marital versus non-marital early childbearing. The percentage of women who had a marital birth by age 20 fell from 14 percent to 3 percent over this period. The comparable statistic for non-marital births rose from 8 percent to 14 percent. Understanding these patterns (including a discussion of whether a decline in shotgun marriages is responsible) is clearly important in understanding the trends in early childbearing.

In Figure 6, we focus on the outcomes for unmarried women who were pregnant before the age of 20. The most notable feature of this figure is the dramatic increase in the percentage of women who go on to carry their pregnancy to term. Although this outcome occurred only about 40 percent of the time in the mid-1970s, it now occurs nearly two-thirds of the time. This increase can be attributed partly to a reduction in the fraction of non-marital conceptions that result in a marital birth (i.e. shotgun marriage). Since around 1980, however, a reduction in the percentage of pregnancy failures, likely the result of less frequent use of abortion, is another contributing factor. In the most recent statistics, a woman who gets pregnant outside of marriage has a 62 percent probability of having a non-marital birth, a 10 percent probability of getting married before the birth, and a 28 percent probability of aborting or having a miscarriage.

B. FFS Data

The Family and Fertility Survey is a dataset that combines survey data from 23 countries mainly in Western and Eastern Europe (along with a few other developed countries) conducted largely during the early and mid-1990s. A standard questionnaire was prepared that asked respondents to report characteristics of their household, parents, partnerships, and children, among other things. Importantly each survey also contains complete fertility histories for each respondent. The survey was given to national representative samples of women (and men in

many countries) of childbearing age. Although a standard questionnaire existed, modifications were imposed in many countries so that the data available is not necessarily uniform.

For our purposes, we restricted our attention to those countries outside of Eastern Europe. Women's fertility histories are the focus of our analysis and childbearing outcomes for women in these FFS countries mostly took place prior to the collapse of the Soviet Union. Because of the vastly different economic environment that existed in those countries during that period, we did not include them in our analysis. The remaining countries include Austria, Belgium, Canada, Finland, France, Germany, Greece, Italy, New Zealand, Norway, Portugal, Spain, and the United States.[9] The United States data actually is the 1995 National Survey of Family Growth. In much of our reported results, we exclude the United States because it is such an outlier among this group of countries in terms of both its level of inequality and the level of early childbearing. As in the NSFG, and for similar reasons, we restrict our attention to those women who were older than 20 or 25 (depending on the age at which outcomes are measured) on the survey date and who turned those ages no earlier than 1976. That date pushes most childbearing outcomes beyond the introduction and diffusion of the birth control pill and, in many (but not all) countries past the legalization of abortion.[10]

Unlike our analysis of NSFG data, we do not distinguish fertility outcomes between those that take place within and outside the scope of marriage. The reason for this is that the relationship between marriage and fertility in many European countries is considerably weaker than it is in the United States; it is not uncommon for committed partners to have children before marriage and then marry sometime later (cf. Kiernan, 2004). As such, the link between

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[9] Obtaining data from the Netherlands requires a separate application procedure, which we have yet to complete.
[10] In our regression models using these data we also include measures of the legal status of abortion in each country in each year. The data necessary to code this variable are available in Levine (2004).

subsequent well-being and whether marriage preceded a birth is weaker, so we do not use marriage to distinguish between outcomes.

To measure a woman's socioeconomic status, we use an indicator of whether a woman grew up in a household ("most of the time") with both of her parents as opposed to a single parent or no parent household. This variable is not available in each country and further restricts the data available to us. In the end, we are left with Austria, Belgium, Finland, Germany, Greece, Italy, Portugal, and Spain as the countries available for our analysis of fertility. When we focus on conception and pregnancy failure, we are further restricted because the surveys in Austria and Germany do not ask about these outcomes, so we are left with data for six countries.

The main additional variable that we attach to these data is the Gini coefficient as our inequality measure that we also calculate as the long-term average within each country. We use the Gini coefficient here rather than the 50/10 ratio because its use in international inequality statistics is more prevalent. We take advantage of the data collection efforts conducted by the United Nations World Institute for Development Economics Research (UN-WIDER), which has cataloged an expansive collection of Gini coefficient estimates for a large number of countries in its World Income Inequality Database.[11] We restrict our attention to all available estimates between 1976 and 2000, focusing on those that are: (a) obtained from nationally representative data sources; (b) deemed to be of high quality; (c) cover the entire population of the country; (d) use individuals as the unit of analysis, and; (e) focus on disposable income. Imposing these restrictions still leaves at least six estimates of the Gini coefficient within each country.[12] We take the simple average of these within-country estimates to obtain our desired long-term

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[11] We obtained these data from http://www.wider.unu.edu/research/Database/en_GB/database/ (accessed June 7, 2011)

[12] For Finland, we have 40 estimates of the Gini coefficient over the 25 years because multiple data sources are available for the same year.

measure of inequality. Along with the Gini coefficient, we also attach national unemployment rates for each country and year along with indicators for the legal status of abortion.

Figure 7 displays trends in conceptions, births, and pregnancy failures in the countries with complete data on all of these outcomes over time. It shows that the percentage of women giving birth by age 20 fell in these countries from about 14 percent to 10 or 11 percent for birth cohorts hitting age 20 in 1976 through 1989 (using this cut-off to maintain the same countries in the panel throughout the period). These rates are at least 50 percent lower than that observed in the United States, as reported in Figure 5. Rates of conception have fallen as well, particularly throughout the 1980s, which is consistent with the drop off in the rate of pregnancy failure during this period. Since changes over time in pregnancy failures are largely attributable to changes in the rate of abortion, the declining occurrence of abortion during this period in Europe matches that occurring in the United States at that time.

Table 2 characterizes the degree of income inequality in all of the FFS countries used in this analysis (including those with missing pregnancy resolution data). It shows a great deal of dispersion in inequality, as captured by the Gini coefficient. As a rough characterization, low inequality countries have a Gini coefficient around .25, middle inequality is characterized by a Gini of around .3 and high inequality is captured by a Gini of around .35. Although not listed here, the estimate of .38 in the United States is higher than any of these other countries. The spread of around .1 is a useful reference point for subsequent interpretation of the magnitude of our regression results later in the paper.

V. RESULTS

A. Analysis of NSFG Data

Before presenting our formal econometric results, we begin by presenting a descriptive analysis of the NSFG data that is comparable in spirit to our regression models that can illustrate the findings to come. To do so, we distinguish states by their long-term level of income inequality (50/10 ratio). As listed in Table 1, we categorize states by quartile, focusing on the top and bottom quartiles along with the interquartile range for each of these income measures, respectively, giving us three categories: high, medium, and low inequality. Then, we distinguish women by their mother's level of education and estimate a set of outcomes for women separately by socioeconomic status (as proxied for by her mother's education group) and state inequality category (high/medium/low). A comparison of outcomes across states and SES groups represents a stylized version of the interaction of SES and state inequality measures, previewing the regressions to follow.

Figures 8A through 8D present the results of this exercise. Figure 8A focuses on non-marital childbearing by age 20. It shows that high SES women (with college-educated mothers) exhibit little variation in early non-marital childbearing across states that differ in their level of income inequality. Low SES women (with mothers who have dropped out of high school) are more likely to give birth to a child at a young age outside of marriage if they live in a high inequality state. Moving from a low inequality state to a high inequality state (which represents roughly a one point increase in the 50/10 ratio) appears to increase the rate of non-marital childbearing by age 20 by around 5 percentage points.

Figures 8B through 8D are designed to determine the antecedent behavior that leads to this increase in non-marital childbearing among low SES women. In Figure 8B we see little evidence that high inequality changes rates of non-marital conceptions differentially for women who differ by SES. Figure 8C displays rates of pregnancy failure that does show a pattern of

behavior across states and groups of women that is consistent with the pattern in non-marital childbearing. Middle and high SES women have no clear pattern in their rates of pregnancy failure, but the pattern for low SES women is clear. Rates of pregnancy failure are a little over 4 percentage points lower for these women when they reside in high inequality states compared to low inequality states. This means that the differences in early non-marital childbearing can almost entirely be attributed to differences in the rate of pregnancy failure, which is likely due to differences in the use of abortion. Figure 8D focuses on rates of shotgun marriage; no apparent pattern is observed across groups here.

These charts represent a stylized version of the regressions reported in Table 3, which reports estimates from our econometric model described earlier. The first column in the upper panel of this table explores non-marital births by age 20 as the outcome. The interaction between the 50/10 ratio and having a high school dropout mother represents β_1 in our econometric model. It indicates that low SES mothers in a high inequality state are more likely to have a non-marital birth by age 20. As we saw earlier, a one point change in the 50/10 ratio roughly captures the movement from a low inequality state to a high inequality state. In this case, that movement is predicted to increase early, non-marital childbearing by 5.3 percentage points for those women whose mothers were high school dropouts. This point estimate is very similar to what we observed in Figure 8A, suggesting that the other covariates in the model have very little correlation with inequality and/or early non-marital childbearing.

The remainder of the top panel of the table focuses on other non-marital outcomes and all marital outcomes by age 20. In Column 3, we see that much of the reason why non-marital childbearing among low SES women rises with inequality is that abortion rates, as captured by pregnancy failures, fall. The magnitude of this estimate indicates that moving from a low

inequality state to a high inequality state reduces the likelihood of a pregnancy failure by 4.2 percentage points. We cannot statistically distinguish this estimate from the 5.3 percentage point increase in early non-marital childbearing. This suggests that the rise in early non-marital births associated with greater inequality is mainly attributable to fewer abortions. Column 2 provides no evidence that the likelihood of contraception is affected. We find little support for an impact of inequality on marital fertility. In Column 7, we focus on shotgun marriage as an outcome and find no statistically significant impact of inequality there for low SES women.

Interestingly, when we focus on moderate SES women as captured by daughters of high school graduates, we find that the point estimate for the increase in non-marital births is attenuated, and now there is a statistically significant reduction in marital births, of nearly the same (opposite-signed) magnitude. Presumably these women are at some, albeit reduced, risk of poor economic outcomes that may be exaggerated in high inequality states. For them, greater inequality is associated with fewer marital births. This suggests that a reduced prevalence of shotgun marriage is the pathway, as confirmed in Column 7. These results are similar to the results for women in their young 20s, described below.

The lower panel of Table 3 replicates this analysis, focusing on childbearing/marital outcomes by age 25 rather than age 20. Including somewhat older young adults increases the relevance of shotgun marriages in these findings. We continue to find that non-marital childbearing increases for low SES women when they live in high inequality states, but we find a similar drop in marital childbearing (although the latter is not quite significant). Changes in the likelihood of a shotgun marriage explain the divergent pattern.[13]

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[13] We have estimated these models separately for blacks and Hispanics, but sample sizes drop substantially and statistical power falls as a result. For the outcomes by age 20, among the sample of Hispanic women (n=4,086), the estimated coefficient (standard error) on the interaction term of interest is 0.060 (.047) for non-marital births, 0.020 (0.050) for non-marital pregnancy failure, and -0.065 (.024) for shot-gun married. These results suggest that low-

B. Analysis of FFS Data

Our conceptual approach in analyzing the FFS data is very similar to that using the NSFG with a few minor distinctions. First, our measure of socioeconomic status is household composition during childhood rather than maternal education. Low SES is determined by whether a woman grew up in a household headed by a single or no parent. Second, we also rely on a country's long-term average Gini coefficient rather than the 50/10 ratio. Third, we no longer make distinctions between marital and non-marital outcomes because this difference has less significance in the European context.

As in our discussion of the NSFG results, we begin by presenting in Figure 9 a graphical depiction of the difference in the likelihood of giving birth by age 20 by socioeconomic status in countries that differ by their level of inequality. Countries are distinguished into inequality categories consistent with the statistics reported in Table 2. Among women who grew up in higher SES households (i.e. with both parents), we see perhaps a trivially small increase in rates of early childbearing among women growing up in higher inequality countries. For women from low SES households, however, a clear pattern is evident that women from high inequality countries are considerably more likely to give birth by age 20. Moving from a low inequality country to a high inequality country increases the odds of having an early birth by around 5 percentage points. Notably, the magnitude of this effect is very similar to that obtained in the U.S. in moving between a low inequality and high inequality state.

SES Hispanic teens living in more unequal places are less likely to get married conditional on getting pregnant. For the outcomes by age 20, among the sample of black women (n=6,117), the estimated coefficient (standard error) on the interaction term of interest is 0.025 (.033) for non-marital births,-0.085 (.025) for non-marital pregnancy failure, and 0.018 (.014) for shot-gun married. These results suggest that low-SES black teens living in more unequal places are less likely to terminate a pregnancy, conditional on getting pregnant. The pattern of results is qualitatively similar for these racial/ethnic groups for outcomes by age 25.

These stylized results are replicated in a more formal econometric analysis, which is reported in Table 4. It reports the results of our econometric model, specified earlier, where β_1 is captured by the interaction between the Gini coefficient and whether a woman was not raised in a two parent household. The top panel considers fertility outcomes by age 20 and the bottom panel by age 25. In interpreting our findings, we focus on the impact of a 0.1 point increase in the Gini coefficient, which roughly reflects a movement from a low inequality country to a high inequality country, as described earlier.

Consistent with Figure 9, the results in the top panel indicate a strong relationship between inequality and early childbearing among low SES women. In Column 1, we present regression results in models that include the United States. In that model, we see that a 0.1 point increase in the Gini coefficient increases the rate of childbearing among low SES women by age 20 by 5.9 percentage points.[14] In Column 2 we drop the United States from the analysis and this estimated impact falls to 3 percentage points. Again, this is consistent with the U.S. being a positive outlier in both inequality and early childbearing along with low rates of two parent families. This estimate falls again, to 2 percentage points, when we restrict our sample in Column 3 to the remaining countries that also report data on other pregnancy outcomes. When we focus on conceptions and pregnancy failures in Columns 4 and 5, respectively, we see that higher inequality generates both more conceptions and fewer pregnancy failures among low SES women by age 20. The importance of conceptions is somewhat different than in the United States, where abortion was the primary determinant.

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[14] Standard errors reported in Table 4 are clustered at the country level. As reported in Cameron, Miller, and Gelbach (2008), even these adjusted standard errors may be understated with so few countries used in the analysis (9, 8, and 6, respectively, in columns 1, 2, and 3 through 5). A simple solution to help address the problem that they describe, which has been implemented in Cohen and Dupas (2010), is to adjust the critical values using a t-distribution with G-2 degrees of freedom, where G is the number of countries. Implementing this approach would change the standard critical value of 1.96 at the 5 percent level of significance from a normal distribution to 2.262, 2.306, and 2.447, respectively, from t-distributions.

The bottom panel of the table focuses on outcomes by age 25. The impact on births is similar to that by age 20, albeit somewhat less precisely estimated. On the other hand, the point estimates in models of conception and pregnancy failure suggest that abortion is a less common mechanism than it is for younger women. The precision of these results, however, is weak enough to prevent us from drawing strong conclusions here.

C. An Investigation of Alternative Mechanisms

Our approach to statistical identification focuses on the interaction between socioeconomic status and a measure of income inequality. That inequality measure is estimated as a long-term average and fixed within states/countries over time. As such, any state/country fixed factor that is highly correlated with the measure of inequality that we use will generate results that are similar in spirit to those reported here. Low SES women in states or countries with that characteristic will be found to have a higher propensity to give birth at an early age. In other words, our interpretation of the role played by inequality may not be warranted if there is some other important state/country fixed factor that is omitted, but strongly correlated with inequality and directly related to childbearing outcomes for low SES women.

Although we do not have a perfect solution to rule out this possibility, we are able to examine the impact of including other plausible alternatives into our model. If including interactions between SES and these other factors reduces the magnitude of the inequality interaction, then this would indicate the presence of omitted variable bias. If not, it would bolster our argument that the relationship between inequality and early childbearing is causal.

In our analysis, we focus on data from the NSFG and the relationship between early (by age 20) non-marital childbearing and other state characteristics interacted with low socioeconomic status. Our emphasis on the NSFG data, as opposed to the FFS, and alternative

state characteristics is more a function of data availability and our ability to generate plausible alternatives than any other substantive consideration.

We consider two sets of alternative factors. The first consists of other features of state income and wage distributions that may matter, but are different than our measure of inequality at the bottom of the income distribution. We consider five alternatives. First, we include the 90/50 ratio from the income distribution, calculated in the same way we described earlier regarding the 50/10 ratio. Since that measure identifies inequality at the top of the distribution rather than the bottom, we would not expect it to have a very strong impact on early childbearing among low SES women. Second (and third), we consider the absolute level of income at the bottom and middle of the income distribution, as opposed to the relative distance between these measures. Perhaps it is not inequality that matters for a young woman's decision, but rather how low the 10th percentile of the income distribution is, or where the 50th percentile is. Fourth (and fifth), we consider the average wage ratio for high school graduates relative to high school dropouts and for college graduates relative to high school graduates. These outcomes provide a measure of the returns to education in the state and provide a way to examine whether a higher return to moving up the ladder can actually reduce the likelihood of early, non-marital childbearing among low SES women.[15]

In Table 5, we report the results of estimating equation [2] including these five alternative measures. Column (1) reports the main coefficients of interest from equation [1] – namely, a point estimate of 0.053 (.015) on the interaction of the 50/10 ratio and our maternal education measure of low-SES status. As we show in the remainder of the table, the coefficients on each of

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[15] We calculated these wage ratios using data the 1980, 1990, and 2000 Censuses along with the 2006-2008 ACS, estimating hourly wages for full-time workers (greater than 1,500 hours in the year) by level of educational attainment.

these alternative measures is not statistically significant (at the 5 percent level) and including

these additional variables, does not attenuate the coefficient on our main variable of interest.

We next consider a set of six alternative state characteristics that are unrelated to

economic outcomes, but capture other dimensions of the political and social climate within a

state. We first consider the political culture of a state, which we capture by including the

average percentage of voters favoring Democratic candidates averaged within state between the

1972 and 2008 Presidential elections.[16] Religious beliefs in a state may also be a contributing

factor; we experiment with including an index of religiosity as well.[17] We also explore the role

played by differences in states' percentage of the population that is minority (defined as

individuals who are not white, non-Hispanic). Next, we include the incarceration rate and the

poverty rate in the state.[18] The final alternative state characteristic is "social capital." Putnam

(2000) argues that the loss of social capital has contributed to the high teen birth rate, so we

include his measure of social capital as a potential alternative.

The results of this analysis are reported in Table 6. Again, the first column replicates the

first column of Table 3 for the purposes of comparison. The results provide no evidence that any

of these additional measures has an effect on early non-marital childbearing rates among low-

SES women. Nor does the inclusion of these variables alter our conclusion about the role of

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[16] The source of these data is David Leip's Atlas of Presidential Elections, available at http://uselectionatlas.org/.
[17] The source of these data is Gallup Poll, *State of the States: Importance of Religion*. January 28th, 2009. Available at: http://www.gallup.com/poll/114022/state-states-importance-religion.aspx.
[18] The incarceration data are compiled by the U.S. Bureau of Justice Statistics, Office of Justice programs, downloaded from www.ojp.usdoj.gov (accessed April, 2011). Poverty rate data comes from the United States Census Bureau and were obtained from http://www.census.gov/hhes/www/poverty/data/historical/people.html (accessed June 7, 2011). In our analysis, we follow the Census Bureau recommendation to use data beginning in 1980 to maintain the consistency of the data over the time period.

inequality.[19] The bulk of the evidence in Tables 5 and 6 is strongly suggestive that inequality has a causal impact on early non-marital childbearing in the United States.

D. Evaluating the Magnitude of the Effects

An important motivation for this current project is the inability of past research to explain much of the geographic variation that exists in teen childbearing. Our analysis suggests income inequality may be an important determinant. The proposed mechanism for the relationship is that greater levels of income inequality lead to a heightened sense of economic despair. We formalize this idea in a model below. Although we can by no means explain all of the variability or even the majority of it, our results are able to explain a sizeable share of the problem. In the United States, according to our estimates 13.6 percent of women experienced a teen, non-marital childbirth in high inequality states compared to 10.1 percent of women in low inequality states over the post-1976 sample window we consider. In those high inequality states, roughly 34 percent of teens were daughters of high school dropout mothers (per our tabulations of the NSFG data). Based on our estimates in Table 3, low SES women in high inequality states were 5.3 percentage points more likely to have a non-marital teen birth compared to low SES women in low inequality states. If the level of inequality in the high inequality states decreased to the level of the low inequality states, we would expect teen, non-marital childbearing to decline by .053*.34*1 = .018 or 1.8 percentage points. This represents 1.8/(13.7-10.2) = .51 or 51 percent of the gap in the likelihood of non-marital childbearing by teens between high and low inequality states. In other words, equalizing rates of income inequality across states would eliminate around half of the difference in rates of early, non-marital childbearing between high and low inequality states.

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[19] One possible exception to this general finding is that including our political measure has a somewhat more substantial downward impact on the key interaction between the 50/10 ratio and our measure of low-SES in the model for non-marital childbearing by age 25.

We also conduct a similar calculation using the results of our analysis of FFS data, with the goal of applying international estimates to explain the gap in teen fertility between the United States and Europe. Focusing exclusively on the data in the FFS, we estimate that 20.1 percent of women give birth before the age of 20 in the United States and 9.2 percent of women in non-Eastern Bloc countries had a teen birth, generating a 10.9 percentage point difference between the U.S. and the other countries. In the United States, 15 percent of women grew up in households headed by a single or no parent. Our estimates in Table 4 that include the U.S. (Column 1) indicate that a one point increase in the Gini coefficient would reduce the likelihood of a birth by age 20 among low SES women by 0.589 percentage points. This means that if those 15 percent of women grew up in two parent families and with the level of inequality observed in a low inequality country with a Gini coefficient of 0.25 instead of the 0.38 value that exists in the U.S., teen fertility in the U.S. would decline by 0.15*(0.25 - 0.38)*0.589 = -0.011. This represents -0.011/0.109 = -0.101, or a 10.1 percent reduction in the gap in teen fertility between the United States and the other countries.

Our estimates suggest that inequality can explain roughly 10 percent of the variation in teen childbearing across countries and 50 percent across U.S. states; however, we suspect that our estimated figure for the cross-state variation is an overestimate and the figure for cross-country variation is an underestimate of the role played by inequality. First, our measure of low-SES status in the cross-country FFS data captures only 15 percent of the population, as compared to our measure of low-SES in the U.S. context, which captures 35 percent of the population. In the FFS, we are probably applying our estimated effect to too small a percent, thereby leading to an artificially low estimate of 10 percent of the gap being explainable by inequality. In addition, the difference in teen birth rates between high and low inequality states is 3 percentage points, as

compared to a 10 percentage point difference between high and low inequality countries. In other words, as a mechanical matter, the denominator of the difference is substantially lower in the cross-state context. Recall from Figures 3 and 4 that the cross-sectional link between inequality and early childbearing is much stronger internationally than across states. Thus, the set of high inequality states are a much more diverse set of states than are the set of high inequality countries. This likely suggests that our 50 percent estimate overstates the amount of cross-state variation that can be explained by inequality. We thus conclude that inequality can explain between 10 and 50 percent of the geographic variation in teen fertility.

VI. AN ECONOMIC MODEL OF THE "CULTURE OF DESPAIR"

To this point, we have empirically addressed the role of inequality in explaining geographic variation in early, non-marital childbearing, but we have not addressed why inequality might matter. Although we do not presume to have a concrete answer to this question, we propose an economic model incorporating insights from the sociological and ethnographic literatures that we described earlier. One of the main contributions of this paper is to show how these ideas are easily captured in a fairly standard economics model of individual decision-making. The model we propose is based on the idea that young girls make decisions about their behaviors – sexual activity, contraceptive use, abortion – based in part on their own perceptions of their likelihood of future success. The primary role of inequality in this model is to affect one's perception of economic success.

In this model, a young, unmarried woman's decision to have a baby is based on a comparison of her expected lifetime utility if she has a baby in the current period compared to expected lifetime utility if she delays childbearing. An individual chooses to have a baby in the current period if the following condition is met:

$$u_o^b + E(V^b) > u_o^d + E(V^d)$$
[3]

where u_o^b is current period utility if she has a baby and u_0^d is current period utility if she delays childbearing. V is the present discounted sum of future period utility.

For young, unmarried women, childbearing has a direct effect on current period utility and an indirect effect on future period utility. We propose that for young, unmarried women of low socioeconomic status (SES) having a baby is utility-enhancing in the current period, such that $u^b > u^d$. This proposition reflects the description from Edin and Kefalas (2005) above, whereby a baby is seen as bringing "purpose, the validation, the companionship, and the order" otherwise missing from many of these women's lives. If $u^b < u^d$, it is never optimal to have a baby in the current period and the model trivially predicts "delay" to the optimizing choice. It seems reasonable to expect that for the majority of high-SES young women, $u^b < u^d$, and the results of the empirical analysis are consistent with that supposition.

For unmarried young women, having a baby in the current period negatively affects expected future utility by leading to lower levels of consumption in the future. For simplicity, we characterize utility in future periods as taking high and low values, U^{high} and U^{low}, respectively. We assume that childbearing at an early age reduces the likelihood of achieving U^{high}. There are two likely mechanisms, the first through the labor market and the second through the marriage market. With regard to the first, we posit that having a baby makes it more difficult for women to acquire human capital, decreasing the future stream of own earnings, and thereby lowering subsequent income and consumption. Having a baby while young and unmarried is also likely to be a hindrance in the marriage market, and would thereby reduce the likelihood of improving one's economic condition through a successful marriage. We define U^{low} to be the level achieved by a young woman who does not delay childbearing. The present discounted value of the young

mother's future utility stream is thus deterministic and captured by V^{low}. If the young woman delays childbearing, there is some positive probability p that she will achieve the "high" utility position in future periods. As we have defined things, a young woman who has a baby in the current period is necessarily assigned to a low position in the income distribution. Our model assumes that if she delays childbearing, she has some probability of achieving the high income/consumption level.[20]

We can therefore write the condition to have a baby in the current period as follows:

$$u_o^b + V^{low} > u_o^d + pV^{high} + (1-p)V^{low}. \tag{4}$$

This condition makes it clear that the change in lifetime utility from delayed childbearing comes from two opposite-signed sources: (1) the loss of current period enjoyment of a baby and (2) a positive probability of achieving the high- utility state in the future. We have implicitly assumed that the delay in childbearing causes no first-order change in the future lifetime enjoyment of the child itself (say, by making childlessness a more likely outcome). In other words, the decision we are modeling is to have a baby in the current period versus having a baby in the subsequent period. So the direct utility loss from not having a child in the current period is limited to the loss in current period utility.

Rearranging terms, we see that a young woman will choose to delay childbearing in the current period if and only if:

$$pV^{high} + (1-p)V^{low} > V^{low} + \left(u_o^b - u_o^d\right) \tag{5}$$

Of course, the young woman does not perfectly observe p. Instead, she bases her decision on her

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[20] Alternatively, we could define "low" and "high" utility as relative constructs that need not correspond to low and high levels in the unconditional income distribution. Defining the positions in the simplest case as corresponding to "low" and "high" positions in the overall income distribution leads the model to have an ambiguous prediction on the relationship between inequality and early non-marital childbearing, as we show below. It is thus a conservative modeling approach, given our main hypothesis.

perception of p. Let us call this subjective probability q, and rewrite the condition for delaying childbearing:

$$qV^{high} + (1-q)V^{low} > V^{low} + \left(u_o^b - u_o^d \right) \qquad [6]$$

If a young woman *perceives* that she has a sizable chance at achieving economic success -- and thereby capturing V^{high} -- by delaying childbearing, the comparison is more likely to favor the choice "delay." On the other hand, if the young woman *perceives* that even if she delays childbearing her chances of economic success are sufficiently unlikely -- in other words, if q is very low -- then the comparison is more likely to favor having a baby in the current period.[21]

Rearranging expression [6], we can define a reservation subjective probability q^r such that a young woman will choose to delay childbearing if and only if:

$$q \geq q^r = \left(\frac{u_o^b - u_o^d}{V^{high} - V^{low}} \right). \qquad [7]$$

We propose that one's perception of the likelihood of economic success, q, depends on (a) her position in the income distribution, as proxied for by SES status, such that $\frac{dq}{d(SES)} > 0$.[22]

We additionally propose that one's perceived probability of success is a function of the interaction of being low SES and inequality, such that $\frac{dq}{d(ineq)} | (SES = low) < 0$. The further

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[21] We are not the first to hypothesize that a notion of opportunity costs is an important determinant of the decision to have a teen birth (for example, this general idea is contained in Lundberg and Plotnick (1995). However, we are not aware of previous work focused on the perception of future economic success and how inequality potentially shapes that perception.

[22] This supposition finds empirical support in tabulations of data from the 1979 National Longitudinal Survey of Youth (NLSY79). That survey includes questions about expectations of future success and perceived control over one's life, as captured by the Rotter Scale Index. We tabulate these variables by maternal education, which we use as our proxy for SES status (as described below). Among young women whose mothers attended college, 32 percent report a high likelihood of achieving her occupational aspirations; this compares to only 18 percent of young women whose mothers are high school dropouts were optimistic. Similarly, on the Rotter Scale of control over one's life (which ranges from 0 for total control to 16 for no control), the average values for daughters of mothers who attended college was 8.01 compared to 9.24 for daughters of high school dropouts.

down in the income distribution one finds herself and the more inequality that exists, the lower is the perception of economic success q. Note that if income inequality and mobility were positively related, inequality would likely not have such an impact on individual's perceptions. We do not have sufficient data on mobility differences across states or countries to include mobility in our models directly, but we offer some evidence in Figure 10 indicating that mobility and inequality are actually negatively related empirically. This figure plots the intergenerational earnings elasticity – which is inversely related to intergenerational mobility – against the Gini coefficient for a set of nine countries.[23] The two variables are strongly positively related, which indicates a strong *negative* correlation between mobility and inequality. To the extent that individuals at the bottom of the income distribution perceive a sizeable degree of income inequality and a low degree of mobility, they are likely to have a lower subjective q, and a higher likelihood of early childbearing.

We have described the decision facing a young woman as primarily being about giving birth. But as considered in the empirical section, there are multiple decision nodes that lead to a non-marital birth: getting pregnant, carrying the pregnancy to term, and not marrying one's partner in a so-called "shot gun marriage." This last pathway explicitly raises the possibility that a woman's decision is influenced not only by her perceived likelihood of her own economic advancement, but also the likelihood that marrying her partner will bring economic success through his economic achievements. Assuming assortative mating, it is easy to see that conditions that lead a young woman to adjust downward her subjective probability of her own economic success will also lead her to have a relatively low subjective probability associated with the likelihood that her male partner will achieve economic success. So, the greater the inequality (and lower the mobility), the lower would be the perceived likelihood that a male

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[23] A similar figure appears in Wilkinson and Pickett (2009).

partner will bring economic advantages. The results presented in Tables 3 and 4 with regard to shot-gun marriage are consistent with this prediction.

The prediction with regard to inequality is not unambiguously signed within the model. Expression [7] shows that q^r varies inversely with the distance between V^{high} and V^{low}, which in the most simple case, may be thought of as inequality. In fact, one could interpret this difference as a greater return to effort.[24] If so, then greater inequality would lower reservation q, and might thereby lead to less early childbearing.[25] As a result, the theoretical prediction of the model regarding inequality is ambiguous. However, our empirical results suggest that the relationship between inequality and childbearing for low SES women is, in fact, positive. The empirical results are consistent with the idea that on net, inequality generates desperation, not aspiration, among those at the bottom of the distribution.

One important element in this model is that future utility is appropriately discounted. The model does not require any present-biased decision making, also known as quasi-hyberbolic discounting, to explain the choice that favors current period utility. If we add present-biased decision making to the model, it would simply amplify the effect of a lower q and make the decision lean even more heavily in favor of having a baby in the current period.

This last pathway explicitly raises the possibility that a woman's decision is influenced not only by her perceived likelihood of her own economic advancement, but also the likelihood that marrying her partner will bring economic success through his economic achievements. Assuming assortative mating, it is easy to see that conditions that lead a young woman to adjust downward her subjective probability of her own economic success will also lead her to have a relatively low

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[24] We loosely tested this proposition earlier by exploring the impact of changes in the college/high school wage premium.

[25] This need not be the case if we define V^{high} to be the "high" level of income/consumption available to the young woman making the choice to delay childbearing, and allow for that upper bound to be distinct from a high position in the unconditional income distribution.

subjective probability associated with the likelihood that her male partner will achieve economic success. So, the greater the inequality (and lower the mobility), the lower would be the perceived likelihood that a male partner will bring economic advantages. The results presented in Tables 3 and 4 with regard to shot-gun marriage are consistent with this prediction.

The value of this model is a framework within which to interpret the behavioral factors that drive the empirical results. Inequality appears to be positively related to early non-marital childbearing among low-SES women. There are a number of proximate ways this could be realized, but the data show that this difference is driven by a greater tendency to "keep the baby" among low-SES girls in more unequal places. Note that this is consistent with the observation of Clark (1965) – cited above – that "*In the ghetto…There is not the demand for abortion or for surrender of the child that one finds in more privileged communities.*"   To the extent that inequality heightens feelings of economic despair or marginalization, our model provides a rationalization for the empirical finding. It also links our econometric analysis and findings to a vast array of social science research that has largely gone unexplored by economists.

VII. FINAL DISCUSSION

This paper has presented a new set of findings regarding the large, persistent cross-sectional variation in teen childbearing. We conducted econometric analyses on two large, individual-level datasets to determine the extent to which income inequality – and other economic and social conditions -- relates to rates of early non-marital childbearing. We find that women who grew up in low socioeconomic status households are substantially more likely to have an early birth (outside of marriage in the United States) when income is more unequally distributed in their location. As the level of inequality increases, low socioeconomic status women are less likely to abort their pregnancies. We econometrically consider other aggregate-

level variables that might affect an individual's perception of economic success, such as the absolute level of income at the bottom of the distribution and the college-high school wage premium. We also considered aggregate-level variables that might be spuriously correlated with inequality and teen birth rates and thereby confound the interpretation of our primary results, such as the political leanings or religiosity of a place. The data do not support any of these alternative explanations.

The analysis and interpretations are guided by a model of economic despair that is built within the paradigm of a utility maximizing rational actor and based on the insights of anthropological and ethnographic evidence. When a poor young woman perceives that socioeconomic success is unachievable to her, she is more likely to embrace motherhood in her current position, as there is little option value to be gained by delaying the immediate gratification of having a baby. When there is relatively more hope of economic advancement, it is relatively more desirable to delay motherhood and invest in human or social capital. We propose that income inequality heightens any perceived sense of economic despair, and so this decision becomes even more common among poor women in more unequal places. Combining our empirical analyses with our economic model, we suggest that an important factor in generating high rates of early, non-marital childbearing in the United States and, particularly, in some states with the United States is young women's perceived lack of economic opportunity.

We offer the following caveats to the interpretation of our empirical results. First, though our model provides a means to interpret the relationship between inequality and early, non-marital childbearing, it need not be the case that it explains the decision-making process of all individuals. For instance, suppose some critical mass of low-SES individuals responds to inequality in the way posited by our model. Others in that same place might then behave

similarly in response to this established social norm. In such a scenario, our model would be relevant to explaining the geographic patterns we observe in early non-marital childbearing, even though it would not account for the full magnitude of the effect.

Second, we have not investigated the precise channels through which inequality might lead to a greater sense of economic despair among adolescents. Various theories exist for how income inequality, as distinct from absolute income, might affect individual-level behavior. Social scientists, particularly political scientists and sociologists, have emphasized the role of *relative*, as distinct from *absolute* deprivation – in leading to acts of social unrest. In the recent economics literature, Luttmer (2005) has documented that people are less happy when they live around people who are richer than themselves. Watson and McLanahan (2011) present evidence that relative income matters for the marriage decision of low-income men. They interpret their model within the idea of an identity construct. We speculate here that there is an alternative channel at play. To the extent that greater levels of income inequality are associated with increased levels of residential and institutional segregation, individuals at the bottom of the income distribution will be more likely to feel a heightened sense of social marginalization, and hence economic despair.[26] This would link directly to the "social isolation" thesis of Wilson (1987). We think this is a promising area of future research that could lead to important insights into how income inequality affects the lives of the poor.

In conclusion, we have presented robust empirical evidence that income inequality is associated with higher rates of early, non-marital childbearing among economically disadvantaged women. Our results suggest that inequality itself, as opposed to other correlated geographic factors, is a primary driver of this relationship. We have proposed a model of

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[26] Watson (2009) presents evidence that as income inequality has increased over time, cities have become increasingly segregated along income lines.

economic despair that could explain these findings: to the extent that income inequality leads to a heightened sense of economic despair among the poor, it will lead to higher rates of early, non-marital childbearing among those at the bottom of the distribution. This could also be part of the explanation for why high-inequality states and countries see higher rates of a host of "drop out" behaviors, including lower educational attainment and higher rates of crime. We consider this a topic worthy of future research.

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Table 1: Household Inequality Measures across States, 1980-2008

Highest Inequality		Middle Inequality		Lowest Inequality	
State	Ratio 50/10	State	Ratio 50/10	State	Ratio 50/10
DC	5.88	NJ	4.24	AZ	3.81
LA	4.92	NM	4.19	ME	3.79
NY	4.74	MI	4.18	AK	3.78
AL	4.65	WV	4.18	MT	3.78
MA	4.61	NC	4.16	IA	3.74
MS	4.52	CT	4.14	NE	3.72
GA	4.51	AR	4.13	WI	3.72
KY	4.49	OH	4.10	WY	3.71
SC	4.39	CA	4.09	VT	3.61
RI	4.39	OK	4.08	ID	3.60
TN	4.32	VA	4.05	NH	3.58
IL	4.32	PA	4.03	NV	3.57
TX	4.30	MO	4.00	UT	3.41
		MD	3.95		
		CO	3.89		
		WA	3.89		
		DE	3.87		
		MN	3.86		
		HI	3.86		
		FL	3.85		
		KS	3.84		
		ND	3.84		
		SD	3.84		
		OR	3.83		
		IN	3.81		

Notes: The highest and lowest inequality groups are the top and bottom quartiles of states, respectively. The middle inequality group is the middle two quartiles of states. These values are calculated from 1970-2008, using U.S. census and ACS data.

Table 2: Household Inequality Measures across FFS Countries, 1976-2000

Low Inequality		Middle Inequality		High Inequality	
Countries	Gini Coeff.	Countries	Gini Coeff.	Countries	Gini Coeff.
Finland	0.230	Germany[*]	0.296	Spain	0.325
Austria[*]	0.246			Italy	0.328
Belgium	0.275			Greece	0.346
				Portugal	0.367

Notes: Countries are divided by apparent breaks in the levels of the Gini coefficients (0.28 and 0.32). The sample is restricted to non-Eastern bloc FFS countries with data on household composition in childhood and it also excludes the United States (whose Gini coefficient in these data is 0.380). Gini coefficients for each country represent the average values of all reported Gini coefficients available from UNU-WIDER for each country, restricting the years considered to those between 1976 and 2000 that are considered to be of high quality, and that cover the full population of households. Countries marked with an asterisk only have data on births, not other pregnancy outcomes.

Table 3: Impact of Long-Term Inequality on
Marital and Non-Marital Fertility Outcomes by Ages 20 and 25, by Socioeconomic Status
(standard errors in parentheses)

| | Non-Marital Outcomes | | | Marital Outcomes | | | "Shotgun Marriage" |
| | Birth | Conception | Pregnancy Failure | Birth | Conception | Pregnancy Failure | |
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				by Age 20			
50/10 Ratio* Mom HS Dropout	0.053 (0.015)	-0.006 (0.018)	-0.042 (0.015)	-0.026 (0.017)	-0.006 (0.016)	0.003 (0.004)	-0.018 (0.009)
50/10 Ratio* Mom HS Graduate	0.021 (0.012)	-0.013 (0.018)	-0.013 (0.015)	-0.027 (0.010)	-0.004 (0.006)	0.002 (0.003)	-0.013 (0.007)
				by Age 25			
50/10 Ratio* Mom HS Dropout	0.040 (0.013)	-0.039 (0.026)	-0.022 (0.025)	-0.041 (0.025)	0.008 (0.029)	-0.010 (0.006)	-0.050 (0.014)
50/10 Ratio* Mom HS Graduate	0.003 (0.014)	-0.049 (0.018)	-0.020 (0.017)	-0.026 (0.012)	0.002 (0.016)	-0.003 (0.004)	-0.021 (0.010)

Notes: reported standard errors are clustered at the state level. Additional explanatory variables in each regression include maternal educational attainment, current age and age squared, race/ethnicity, an indicator variable for living with a single parent at age 14, the state unemployment rate at age 19, state welfare policies (family cap and maximum AFDC/TANF benefit for a family of 3), state abortion policies (Medicaid funding, parental notification/consent, and mandatory delay laws), and an indicator variable for SCHIP implementation, along with state and cohort fixed effects. The sample sizes are 24,720 and 23,037 in the models by age 20 and age 25, respectively.

Table 4: Impact of Long-Term Inequality on Fertility Outcomes by Ages 20 and 25,
by Socioeconomic Status
(standard errors in parentheses)

	Birth (Full Sample including US) (1)	Birth (Full Sample excluding US) (2)	Birth (restricted sample) (3)	Conception (4)	Pregnancy Failure (5)
			by Age 20		
Gini coefficient* Not Raised in Two Parent HH	0.589 (0.157)	0.301 (0.116)	0.192 (0.048)	0.180 (0.059)	-0.111 (0.015)
Sample Size	29,671	23,042	15,546	15,546	15,546
			by Age 25		
Gini coefficient* Not Raised in Two Parent HH	0.545 (0.181)	0.161 (0.172)	0.249 (0.138)	0.292 (0.163)	-0.010 (0.112)
Sample Size	29,846	22,686	15,509	15,509	15,509

Notes: The full sample includes those women in the FFS outside of the Eastern bloc (with or without the United States, as indicated). The restricted sample includes data from just the subset of non-US countries that also include information about conceptions and pregnancy failures. Standard errors are clustered at the country level (see the text for a discussion of this).

Table 5: Impact of Alternative State Economic Conditions on
Non-Marital Fertility by Ages 20 and 25, by Socioeconomic Status
(standard errors in parentheses)

|	50/10 ratio (1)	90/50 ratio (2)	10th Percentile of Income (3)	50th Percentile of Income (4)	HS Grad to HS Dropout Wage Premium (5)	College Grad to HS Grad Wage Premium (6)
Correlation between 50/10 ratio and characteristic:	---	0.709	-0.615	-0.168	0.127	0.347
by Age 20						
50/10 Ratio* Mom HS Dropout	0.053 (0.015)	0.069 (0.021)	0.062 (0.019)	0.056 (0.015)	0.052 (0.015)	0.054 (0.014)
50/10 Ratio* Mom HS Graduate	0.021 (0.012)	0.000 (0.016)	0.022 (0.017)	0.022 (0.012)	0.017 (0.011)	0.012 (0.011)
State Characteristic* Mom HS Dropout	--- ---	-0.065 (0.066)	0.028 (0.049)	0.028 (0.049)	0.049 (0.093)	-0.035 (0.072)
State Characteristic* Mom HS Graduate	--- ---	0.076 (0.044)	0.003 (0.037)	0.005 (0.037)	0.128 (0.058)	0.097 (0.060)
by Age 25						
50/10 Ratio* Mom HS Dropout	0.040 (0.013)	0.057 (0.018)	0.034 (0.019)	0.039 (0.014)	0.042 (0.013)	0.045 (0.013)
50/10 Ratio* Mom HS Graduate	0.003 (0.014)	-0.006 (0.018)	0.003 (0.019)	0.003 (0.014)	0.003 (0.014)	-0.006 (0.013)
State Characteristic* Mom HS Dropout	--- ---	-0.064 (0.055)	-0.020 (0.049)	-0.021 (0.048)	-0.106 (0.074)	-0.089 (0.083)
State Characteristic* Mom HS Graduate	--- ---	0.030 (0.047)	0.001 (0.047)	0.001 (0.046)	-0.003 (0.059)	0.092 (0.064)

Notes: reported standard errors are clustered at the state level. Additional explanatory variables in each regression include maternal educational attainment, current age and age squared, race/ethnicity, an indicator variable for living with a single parent at age 14, the state unemployment rate at age 19 or 24, state welfare policies (family cap, waiver/TANF implementation, and maximum AFDC/TANF benefit for a family of 3), state abortion policies (Medicaid funding, parental notification/consent, and mandatory delay laws), and an indicator variable for SCHIP implementation, along with state and cohort fixed effects. The specifications for non-marital fertility have a sample size of 24,720 and 23,037 by ages 20 and 25, respectively.

Table 6: Impact of Alternative State Characteristics on
Non-Marital Fertility by Ages 20 and 25, by Socioeconomic Status
(standard errors in parentheses)

|	50/10 ratio (1)	Percentage of Votes to Democrats (2)	Index of Religiosity (3)	Percentage of Population that is Minority (4)	Incarceration Rate (5)	Poverty Rate (6)	Social Capital Index (7)
Correlation between 50/10 ratio and characteristic:	---	0.358	0.249	0.368	0.414	0.499	-0.598
by age 20							
50/10 Ratio* Mom HS Dropout	0.053 (0.015)	0.047 (0.019)	0.057 (0.016)	0.053 (0.016)	0.058 (0.018)	0.060 (0.024)	0.049 (0.016)
50/10 Ratio* Mom HS Graduate	0.021 (0.012)	0.019 (0.015)	0.019 (0.010)	0.010 (0.011)	0.012 (0.012)	0.021 (0.021)	0.015 (0.016)
State Characteristic* Mom HS Dropout	--- ---	0.131 (0.121)	-0.047 (0.079)	0.000 (0.001)	-0.004 (0.006)	0.001 (0.002)	-0.003 (0.012)
State Characteristic* Mom HS Graduate	--- ---	0.038 (0.095)	0.043 (0.061)	0.001 (0.001)	0.009 (0.005)	0.001 (0.002)	-0.007 (0.013)
by Age 25							
50/10 Ratio* Mom HS Dropout	0.040 (0.013)	0.027 (0.015)	0.047 (0.011)	0.048 (0.013)	0.046 (0.012)	0.030 (0.021)	0.049 (0.011)
50/10 Ratio* Mom HS Graduate	0.003 (0.014)	-0.008 (0.014)	0.002 (0.015)	-0.004 (0.014)	-0.004 (0.017)	0.001 (0.021)	-0.016 (0.020)
State Characteristic* Mom HS Dropout	--- ---	0.230 (0.123)	-0.089 (0.074)	-0.001 (0.001)	-0.009 (0.007)	0.001 (0.003)	0.019 (0.015)
State Characteristic* Mom HS Graduate	--- ---	0.163 (0.097)	0.008 (0.062)	0.000 (0.001)	0.003 (0.006)	0.000 (0.002)	-0.013 (0.013)

Notes: reported standard errors are clustered at the state level. Additional explanatory variables in each regression include maternal educational attainment, current age and age squared, race/ethnicity, an indicator variable for living with a single parent at age 14, the state unemployment rate at age 19 or 24, state welfare policies (family cap, waiver/TANF implementation, and maximum AFDC/TANF benefit for a family of 3), state abortion policies (Medicaid funding, parental notification/consent, and mandatory delay laws), and an indicator variable for SCHIP implementation, along with state and cohort fixed effects The specifications for non-marital fertility have a sample size of 24,720 and 23,037 by ages 20 and 25, respectively (and 19,797 and 18,940 for the poverty rate models).

**Figure 1: International Comparison of Teen Birth Rates, 2008**
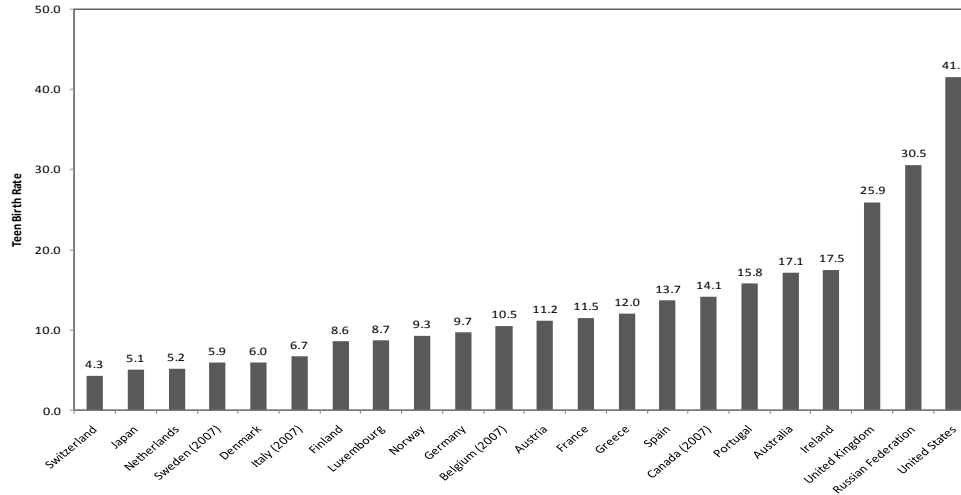
source: United Nationsl Economic Commission for Europe (2011) and UN Demographic Yearbook, 2008.

**Figure 2: Variation in Teen Birth Rates across States, 2008**
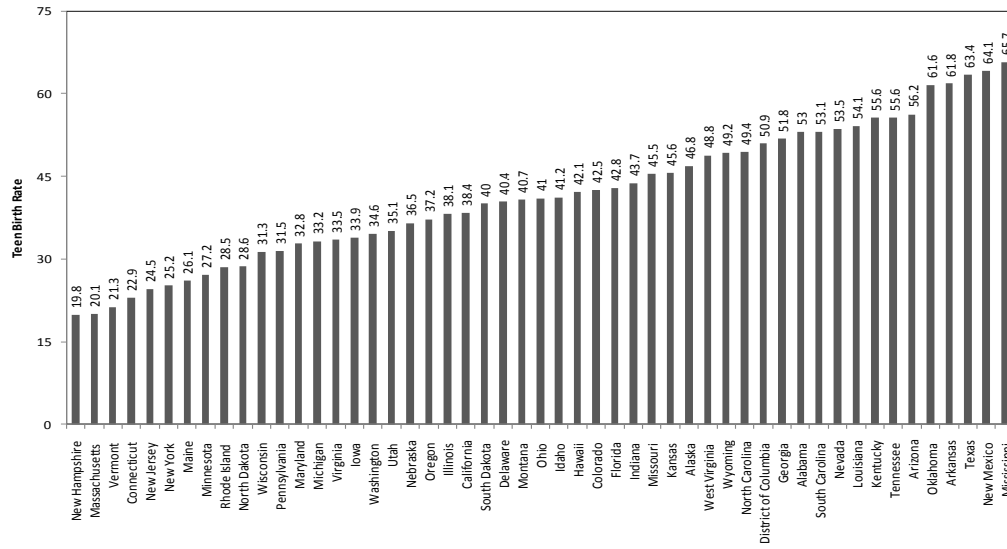
source: Martin, et al. (2010)

Figure 3: Income Inequality and Teen Birth Rates across Countries


United States ◆

$y = 140.17x - 31.554$
$R^2 = 0.502$

Teen Birth Rate (y-axis), Gini Coefficient (x-axis)
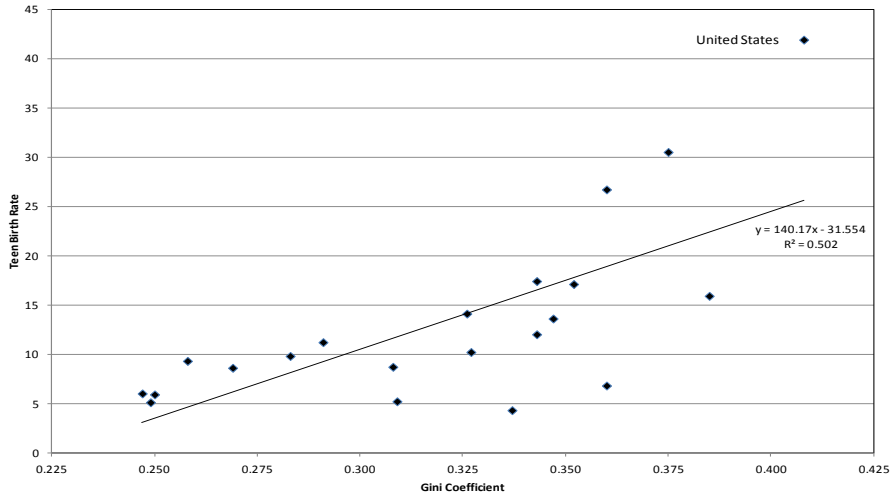
sources: Gini Coefficient - United Nations (2009). Teen Birth Rate - United Nations (2008)

Figure 4: Income Inequality and Teen Birth Rates across States

$y = 236.99x - 64.649$
$R^2 = 0.1313$

CT ◆   NY ◆

Teen Birth Rate (y-axis), Gini Coefficient (x-axis)
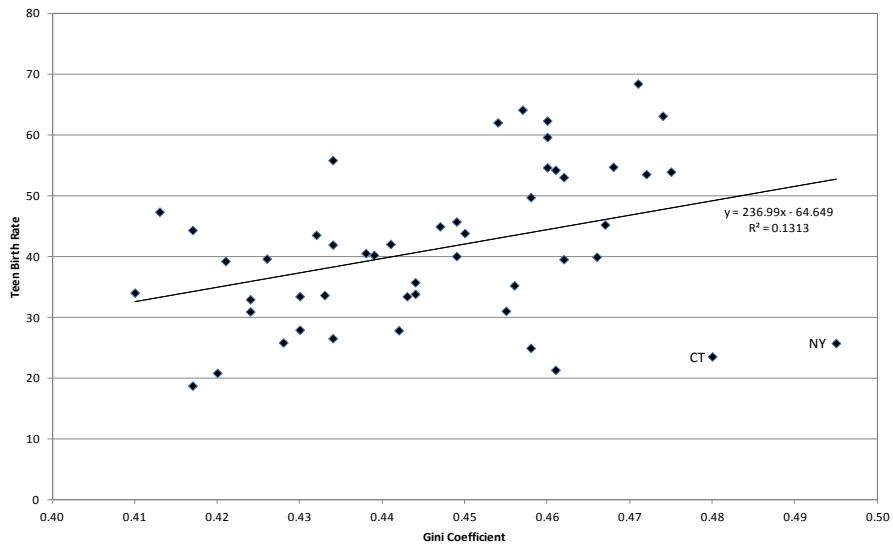
Source: Gini Coefficient - Webster (2007). Teen Birth Rate - Martin (2009)

Figure 5: Probability of Giving Birth by Age 20, by Marital Status
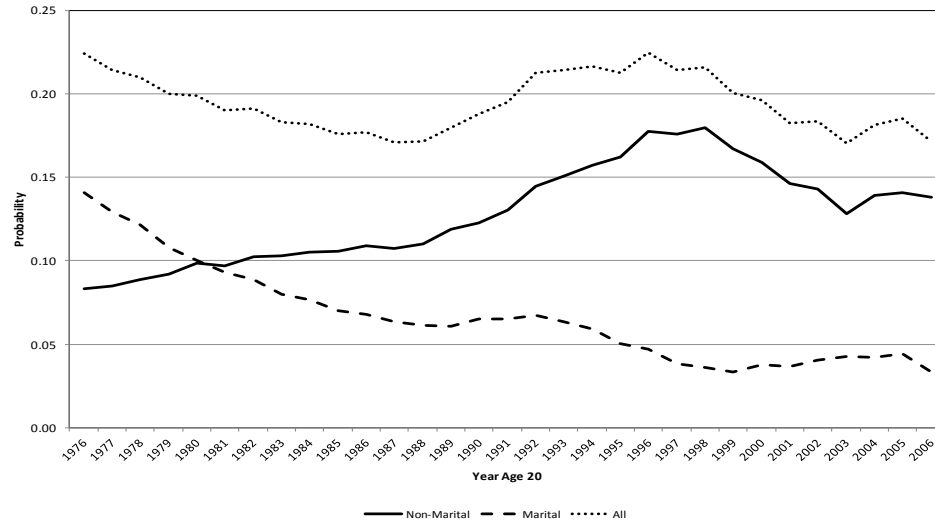
note: Statistics reflect the five year moving average centered on the reported year age 20, weighted by the number of observations.

Figure 6: Pregnancy Resolution for Non-Marital Conceptions by Age 20
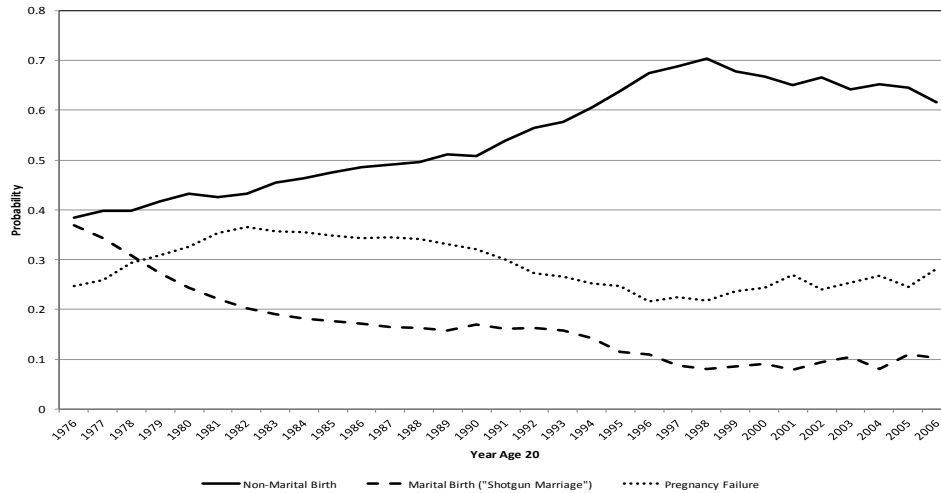
note: Statistics reflect the five year moving average centered on the reported year age 20, weighted by the number of observations.

**Figure 7: Fertility Outcomes by Age 20 in Selected FFS Countries**
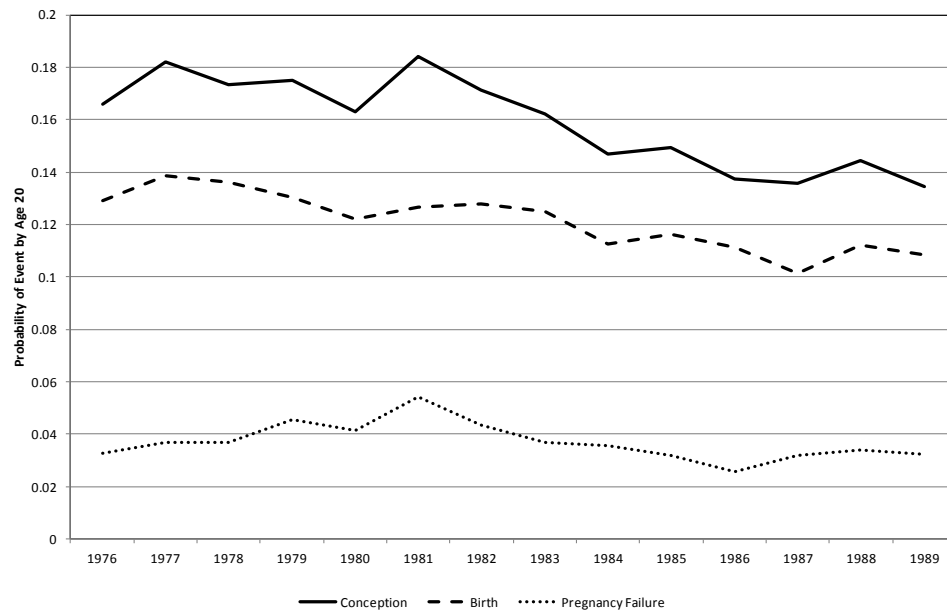
Note: Eastern bloc countries are excluded. Others are included based on the availability of data on all outcomes (see text).

Figure 8A: Rate of Nonmarital Childbearing by Age 20, by Mother's Level of Education and State Level of Income Inequality
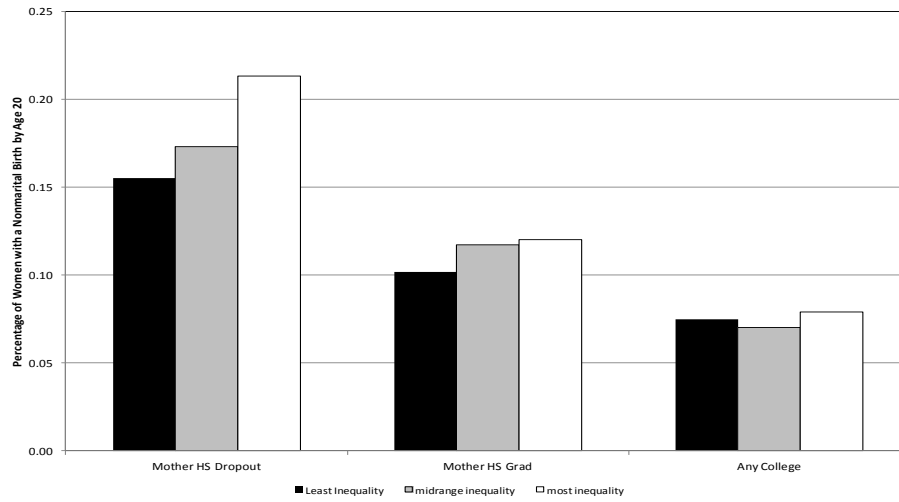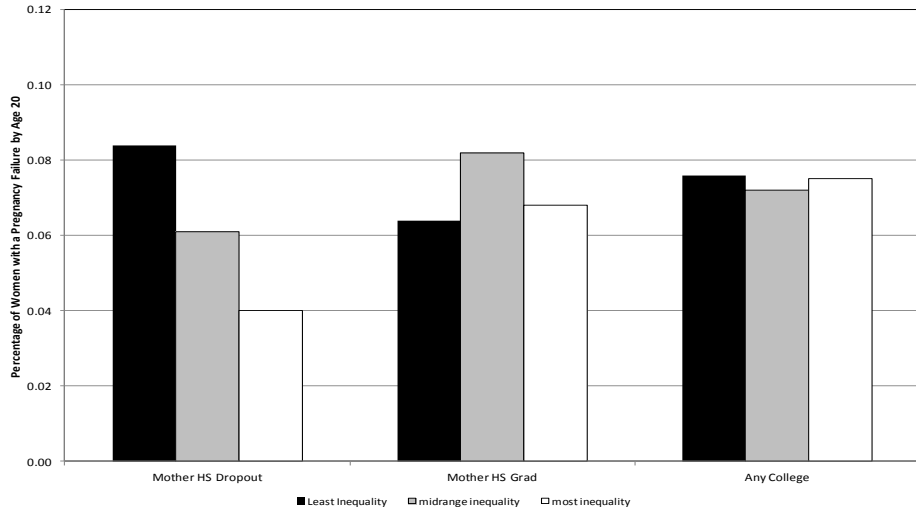

Figure 8B: Rate of Nonmarital Conception by Age 20, by Mother's Level of Education and State Level of Income Inequality
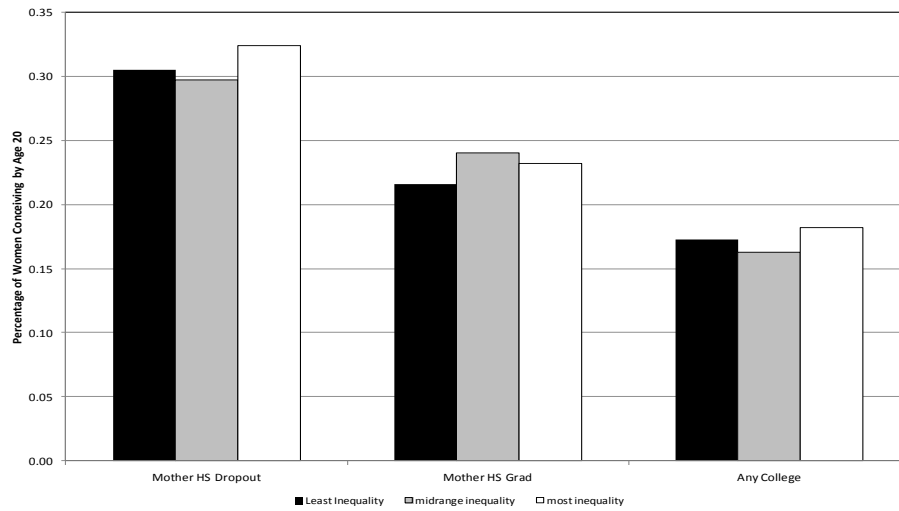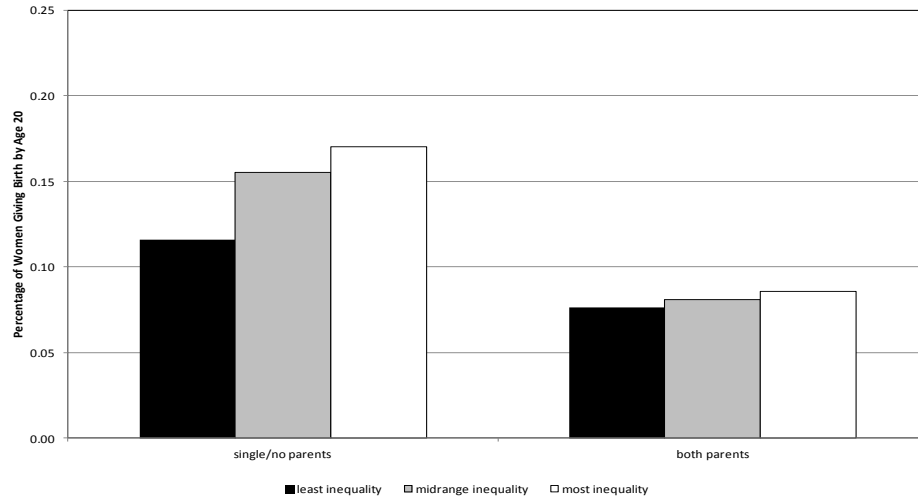
**Figure 8C: Rate of Nonmarital Pregnancy Failure by Age 20,
by Mother's Level of Education and State Level of Income Inequality**
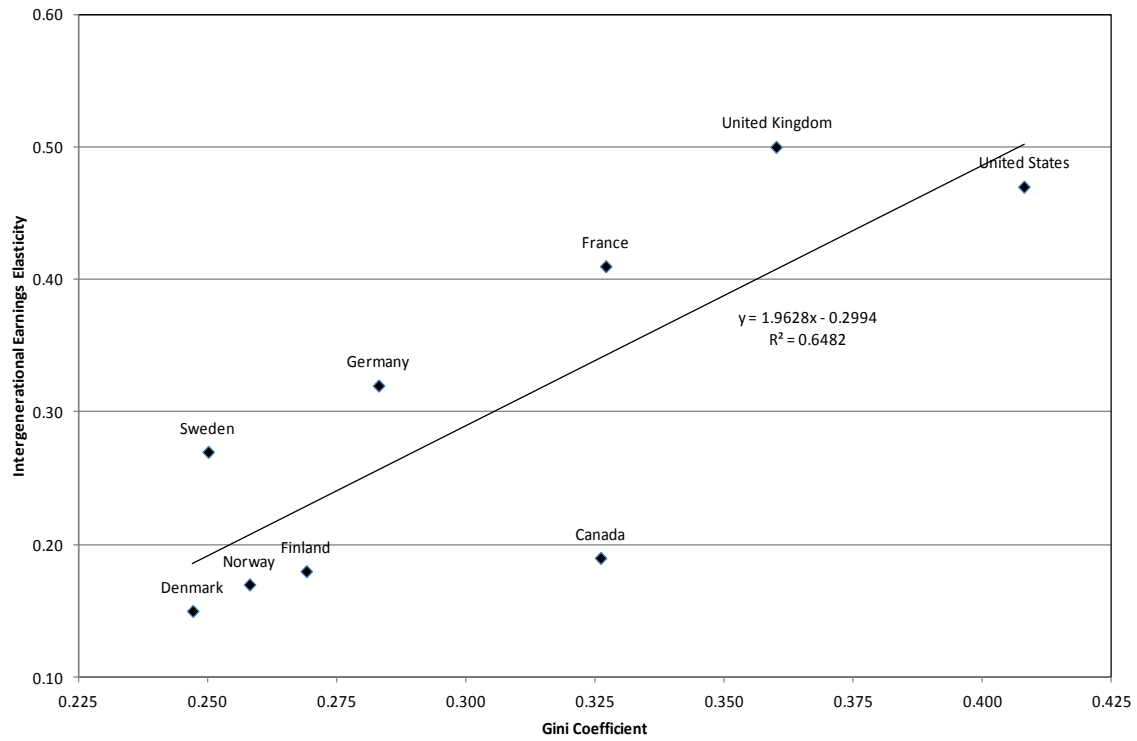

**Figure 9: Rate of Childbearing by Age 20,
by Parental Presence as a Child and National Level of Income Inequality**

source: authors calculations using data from the Fertility and Family Surveys, excluding data from the United States.

Figure 10: Relationship between Income Inequality and Mobility

y = 1.9628x - 0.2994
R² = 0.6482

Axis labels: Intergenerational Earnings Elasticity (y-axis), Gini Coefficient (x-axis)

Data points: United Kingdom, United States, France, Germany, Sweden, Canada, Finland, Norway, Denmark

Sources: Gini Coefficient - United Nations (2009). Intergenerational earnings elasticity - Corak (2006).

Data Appendix: State/Year Policy and Economic Condition Variables

This table lists the sources for the policy and economic condition variables included as control variables in our OLS regressions.

Family cap		Information on this welfare waiver policy was obtained from three sources: (1) a technical report of the Council of Economic Advisers (1999); (2) an Urban Institute report written by Gallagher, et al. (1998); and (3) a report by Crouse (1999), prepared for the U.S. Department of Health and Human Services, which summarizes information contained in a report of the U.S. Department of Health and Human Services (1997). Full source citations: Council of Economic Advisors. 1999. "Technical Report: Economic Expansion, Welfare Reform, and the Decline in Welfare Caseloads: An Update." Washington, D.C.: Executive Office of the President of the United States; Crouse, Gil. 1999. "State Implementation of Major Changes to Welfare Policies, 1992-1998." Washington, D.C.: Office of Human Services Policy, Assistant Secretary for Planning and Evaluation, U.S. Department of Health and Human Services; Gallagher, L. Jerome, Megan Gallagher, Kevin Perese, Susan Schreiber, and Keith Watson. 1998. *One Year after Federal Welfare Reform: A Description of State Temporary Assistance for Needy Families (TANF) Decisions as of October 1997*. Washington D.C.: Urban Institute, Occasional Paper Number 6.
SCHIP implementation		Dates for state implementation of SCHIP were obtained from the Centers for Medicare and Medicaid Services (CMS) website: http://www.cms.hhs.gov/schip/enrollment/fy2000.pdf, accessed September 2005. (All states implemented SCHIP by 1998, so these series did not require updating.)
AFDC/TANF Benefit for 3-person Family		Data on welfare benefit levels were obtained from Council of Economic Advisers (1997) and Rowe, et al., (2004), Rowe and Russell (2004), and Rowe and Versteeg (2005). We update these series using data compiled by the University of Kentucky Center for Poverty Research, publically available on the Center's website. They list their sources as Urban Institute 2006-2009. Full source citations: Council of Economic Advisers. 1997. Technical Report: Explaining the Decline in Welfare Receipt, 1993-1996. Washington, D.C.: Executive Office of the President of the United States; Rowe, Gretchen, Kevin McManus, and Tracy Roberts. 2004. *The Welfare Rules Databook: State Policy as of July 2001*. Washington, D.C.: The Urban Institute; Rowe, Gretchen with Victoria Russell. 2004. *The Welfare Rules Databook: State Policy as of July 2002*. Washington, D.C.: The Urban Institute, October; Rowe, Gretchen with Jeffrey Versteeg. 2005. *The Welfare Rules Databook: State Policy as of July 2003*. Washington, D.C.: The Urban Institute.
Medicaid Family		See Table 1 of Kearney and Levine (2008) for details and descriptions of

Planning Waiver, Income Based Medicaid Family Planning Waiver, Duration Based		these policies through 2006. Sources for those codings: Jennifer J. Frost, Adam Sonfield, and Rebecca Benson Gold. 2006. "Estimating the Impact of Expanding Medicaid Eligibility for Family Planning Services," Occasional Report, New York: Guttmacher Institute, No. 28; and Guttmacher Institute. 2006. "State Medicaid family planning eligibility expansions," *State Policies in Brief*, New York: Guttmacher Institute. We update these series using information reported by recent Guttmacher state policies in brief factsheets, accessed through the website: http://www.guttmacher.org/statecenter/spibs/spib_SMFPE.pdf. We verify this information using information posted on the Center for Medicaid and Medicare (CMS) website about Section 1115 state waivers: http://www.cms.gov/MedicaidStWaivProgDemoPGI/MWDL/list.asp. (Accessed 10/14/10).
Abortion restriction - Parental consent Abortion restriction – mandatory delay		Levine (2004) includes a detailed description of these restrictions and how the variables are coded. We updated Levine's earlier series by comparing changes in legal status between 2004 and what is reported by Guttmacher as 2010 law. For the set of states with reported changes, we searched the state websites for information about dates of implementation: http://prochoiceamerica.org/government-and-you/state-governments/.
Medicaid Abortion State funding restriction		Levine (2004) includes a detailed description of these restrictions and how the variables are coded. We updated Levine's earlier series using information from NARAL, "Restrictions on Low Income Women's Access to Abortion", accessed 10/15/10: http://www.prochoiceamerica.org/what-is-choice/fast-facts/low-income-women.html
state unemployment rate		Bureau of Labor Statistics, Local Area Unemployment Statistics (LAUS) program.