Booms, Busts, and Fertility: Testing the Becker Model Using Gender-Specific Labor Demand

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ABSTRACT

In this paper, I present estimates of the effect of local labor demand shocks on birth rates. To identify exogenous variation in male and female labor demand, I create indices that exploit cross-sectional variation in industry composition, changes in gender-education composition within industries, and growth in national industry employment. Consistent with economic theory, I find that improvements in men's labor market conditions are associated with increases in fertility while improvements in women's labor market conditions have smaller negative effects. I separately find that increases in unemployment rates are associated with small decreases in birth rates at the state level.

I. Introduction

This paper examines the degree to which changes in fertility rates in the United States from 1980–2009 can be attributed to demand-induced changes in labor market opportunities for men and women. Economic theory, originating with the work of Gary Becker (1960, 1965), suggests that positive shocks to family income and improvements in male labor market conditions should be associated with increases in fertility, while better wages and employment opportunities for women should have opposing income and substitution effects. These theoretical predictions have served

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as a foundation for economic analysis of fertility for more than half a century, providing a basis for interpretation of cross-sectional patterns, secular trends, and cyclical fluctuations in fertility. However, despite the prevalence of Becker's fertility model, identification of the model's parameters has proven elusive until recently. Moreover, even with the recent resurgence of interest among economists in identifying the determinants of fertility, there remains little research on the causal relationship between women's labor market opportunities and birth rates and how this relationship has influenced aggregate movements in fertility in the United States over the past several decades.

The primary challenge to empirical tests of theoretical models of fertility has been a lack of exogenous variation in the relevant explanatory variables: family income, male earnings, and female wages. At the individual level, measurement error in income and earnings, unobserved preferences for work and children, and correlation between male and female labor market outcomes are likely to bias estimates. And, while aggregation can alleviate some of these concerns, endogeneity of observed income and wage measures at the aggregate level remains an issue.[1] Recently, researchers have had some success in identifying exogenous shocks to family income and male earnings (Lindo 2010, Black et al. 2012, Lovenheim and Mumford 2013, Dettling and Kearney 2014). However, it has proven more difficult to identify exogenous variation in female labor market conditions. Most studies have relied on variation in observed female wages or unemployment rates (see, for example, Butz and Ward 1979, Macunovich 1995, Adsera 2005, Örsal and Goldstein 2010). These variables are problematic in that they are influenced not only by changes in demand for female labor but also by changes in female labor supply, which in turn is correlated with fertility preferences.[2]

In this paper, I present a new way to identify the causal effects of transitory fluctuations in both male and female labor market conditions on fertility: I create gender-education-specific shift-share indices of labor demand that exploit variation in industry employment shares across states (and sometimes demographic groups), differences in employment growth rates across industries, and changes in the share of men and women of different levels of educational attainment in each industry over time. Building on the instrument used by Bartik (1991), Katz and Murphy (1992), and Blanchard and Katz (1992), among others, this strategy addresses both the measurement error bias and the potential endogeneity that arise when considering gender-specific labor market conditions. Because the industry composition component of the instrument is fixed at a single point in time and the employment growth and gender-share components vary at the national level, the instrument is unlikely to be influenced by local changes in fertility or preferences.

To motivate a more careful look at the relationship between gender-specific labor market conditions and birth rates, I begin by using the shift-share approach to investigate how fertility has responded to changes in general economic conditions in the

1. Hotz, Klerman, and Willis (1997) provide a detailed discussion of these and other sources of bias in empirical estimates of the parameters in theoretical models of fertility.
2. Del Bono, Weber, and Winter-Ebmer (2012) examine the effects of male and female job loss due to plant closure on birth rates. However, as involuntary job loss is a stressful event associated with both permanent decreases in income and involuntary shocks to time use for a nonrepresentative group of individuals, their methodology is arguably unsuitable for identifying substitution effects from transitory changes in labor market opportunities.

United States over the past several decades. The results contribute new estimates to a large literature examining movements in fertility over the business cycle that has yet to come to a consensus.[3] I find that fertility is procyclical—a one percentage point increase in unemployment is associated with a 0.85 to 2.2 percent decrease in the birth rate. Using a shift-share labor demand index as an instrument to address endogeneity and measurement error in unemployment rates increases the magnitude of the estimated relationship between unemployment and fertility. I also explore differences in the response of fertility to local economic conditions by race, age, and educational attainment. I find consistently negative effects for all demographic groups, with IV estimates showing slightly stronger procyclicality in the birth rates of older, white, and more-educated individuals.

Reduced-form analysis using two different types of gender-specific labor demand indices shows that improved labor market conditions for men are associated with increases in fertility for women of all levels of educational attainment. In contrast, the coefficients on female labor demand indices are consistently negative, smaller in magnitude than the estimated male effects, and only sometimes significantly different from zero. I find that the negative effects of improvements in female labor demand are stronger for black women than for white women and are decreasing in magnitude with educational attainment, which suggests that rates of single parenthood and utilization of market-based childcare may be important determinants of the responsiveness to female labor demand shocks within subgroups.

The results in this paper provide both a new baseline estimate of the relationship between economic conditions and fertility in the United States and an empirical test of the relative importance of the income and substitution effects resulting from changes in male income and female wages. The estimated effects are consistent with a key prediction of economic models of fertility: that exogenous improvements in men's labor market conditions should have a positive effect on fertility. Also consistent with economic models of fertility is the result that, holding men's labor-market conditions constant, there is a negative correlation between women's labor market conditions and birth rates. This is among the first papers to provide evidence of substitution effects using a plausibly exogenous source of demand variation.

II. Theory and Literature

A. Theoretical Framework

Economic theory of fertility outlined in early papers by Becker (1960, 1965), Mincer (1963), Becker and Lewis (1973), and Willis (1973) has served as a foundation for empirical analyses of fertility over the past half century, including many studies that

3. While most researchers agree that fertility is procyclical (Becker 1960, Ben-Porath 1973, Macunovich 1995, Andersson 2000, Adsera 2005, Lindo 2013), there are exceptions, including Butz and Ward (1979), and Mocan (1995). Of particular relevance to this paper are the findings of Dehejia and Lleras-Muney (2004), who use a state panel of vital statistics data to study infant health and find no significant relationship between unemployment rates and birth rates despite observing changes in selection into motherhood over the business cycle.

consider the response of fertility rates to variation in economic conditions.[4] In the basic static model of fertility, parents choose the number of children that maximizes their utility, subject to a family budget constraint. The production of children is typically assumed to require parental time in addition to monetary inputs, so that increases in wages are associated with increases in the price of children. Assuming that females are the primary caregivers and that children are normal goods, this model predicts that an increase in family income or male earnings should increase the total demand for children, while an increase in female wages will have both positive income effects and negative substitution effects on fertility.

Though straightforward theoretically, predictions from this simple static model about the causal effects of family income and male and female wages have proven difficult to test empirically. A long-running discussion in the economics literature about how to reconcile the notion of children as a normal consumption good with the negative correlation between income and fertility seen in both cross-sectional and time series data (see, for example, Hotz, Klerman, and Willis 1997; Jones and Tertilt 2008) highlights the importance of identifying exogenous variation in income from different sources. Though the "price of time hypothesis"—the idea that higher income is associated with a higher value of female time and thus a higher price of children—can potentially explain the negative income-fertility relationship, it is likely that the observed data patterns also reflect unobserved heterogeneity in preferences (Jones, Schoonbroodt, and Tertilt 2010). Another theory, first proposed by Becker (1960) in his seminal work, posits that parents have preferences over both child quantity and child quality, with different demand elasticities for each.[5]

Building on the basic static model, dynamic or lifecycle models of fertility assume that individuals can control not only their completed fertility but also the timing of births over their life cycle (Happel, Hill, and Low 1984; Hotz, Klerman, and Willis 1997). In the case of perfect certainty and perfect capital markets, transitory fluctuations in wages do not alter expected lifetime income and thus should not impact expected total fertility. They do, however, impact the timing of fertility if couples respond to transitory fluctuations in female wages by choosing to give birth when wages are low. In the presence of uncertainty and imperfect capital markets, the predictions of the static model hold: Transitory wage changes should have both income and substitution effects, particularly for those groups that are most likely to be credit-constrained.

In addition to differences in access to credit, there are other reasons to expect that the responsiveness of fertility to economic shocks might vary for women of different sociodemographic groups. For example, groups may have differing degrees of knowledge of and access to contraception, which might affect the ability to control the

4. Formal versions of the theoretical models of the determinants of fertility discussed in this paper can be found in Hotz, Klerman, and Willis (1997) and Jones, Schoonbroodt, and Tertilt (2010).

5. Though the quantity/quality tradeoff has been suggested as another possible explanation for the negative income-fertility puzzle (Becker and Lewis 1973, Willis 1973), quality adjustments have recently been viewed more often as a potential substitute for adjustments on the quantity margin. Jones, Schoonbroodt, and Tertilt (2010) offer the following interpretation of Becker's (1960) theory: "Becker argues, by analogy with other durable goods, that economic theory suggests that fertility and income should be positively related, but perhaps only weakly so, while quality of children and income should be strongly positively correlated." Because the emphasis here is on birth rates as an outcome variable, this paper abstracts from further discussion of child quality.

timing of fertility. Differences in access to and use of market-based childcare options may also generate differences in the magnitude of substitution effects for women at different wage levels (Perry 2004; Jones, Schoonbroodt, and Tertilt 2010). In particular, for women whose wages are well above the price of market-based childcare and thus purchase the majority of their childcare, the substitution effect of a wage increase may be weaker relative to the income effect, as the "price" of basic childcare is not directly linked to the woman's own wages.[6] Additionally, single women may experience stronger substitution effects than married women, as they are less likely to share the burden of childcare with a partner.

Because the panel data approach used in this paper exploits short-term fluctuations in state labor market conditions around their long-run trends, and because individuals will be able to adjust to labor market shocks by changing either their industry of employment or their geographic location[7] in the long run, I proceed with the dynamic models of fertility in mind. Thus, I estimate the effects of changes in general and gender-specific labor market conditions[8] on contemporaneous fertility at the aggregate level, and interpret the response as a shift in the timing of fertility across the life cycle. I note, however, that if business cycle fluctuations have permanent income effects on families or if delaying fertility results in lower completed fertility for some women, my results will be picking up not only temporary shifts in the timing of fertility but also changes in total lifetime fertility. Because there are a number of reasons to expect heterogeneity in the responsiveness of fertility to general economic shocks and in the magnitudes of the income and substitution effects from a shock to female wages, I separately estimate the effects of unemployment and male and female labor demand on fertility by race, age, and educational attainment.

B. Related Empirical Work

As economic models of fertility choice emphasize important differences between the effects of changes in male and female wages, predictions from these models about the nature of movements in fertility rates over the business cycle are ambiguous. In particular, because labor market opportunities for both men and women decline in recessions, it is unclear whether income or price effects will dominate. In a highly cited paper entitled "The Emergence of Countercyclical U.S. Fertility," Butz and Ward (1979) hypothesized that the increase in women's labor force participation in the lat-

6. However, if there are inputs into child production that mothers cannot purchase in the market or if leisure time and children are complementary, substitution effects may still exist, even for high-wage women.

7. The desire to minimize the bias from selective migration in response to economic shocks informs the decision to conduct this analysis at the state level and to include controls for state demographic composition. Reassuringly, Wozniak (2010), Lindo (2013), and others find that highly educated individuals and those with relatively high permanent income are more likely to leave an area in response to an economic downturn. Because these groups have lower fertility, it is likely that my approach will underestimate the negative effects of economic downturns on birth rates if the demographic controls do not sufficiently capture the changes in population composition.

8. In this paper, as in much of the literature on the effects of aggregate economic conditions on individual outcomes, terms like "labor market conditions" and "economic downturns" are intentionally vague. This is because aggregate indicators like unemployment and employment rates are not just capturing changes in employment but also all of the other changes to the economy over the business cycle including changes in work hours, wages, and perceptions.

ter half of the 20th century would increase the relative importance of the substitution effect, causing a permanent reversal of the previously positive relationship between economic conditions and birth rates. However, the empirical evidence on the cyclicality of birth rates since the publication of Butz and Ward's paper has been mixed.

Meanwhile, most attempts to directly test the predictions from economic models of fertility using aggregate data have relied on potentially endogenous measures of male and female labor market conditions. To test their counter-cyclicality hypothesis, Butz and Ward include male annual earnings and the female hourly wage in their analysis. Adsera (2005) focuses on the difference between female and male unemployment rates in European countries, while Örsal and Goldstein (2010) examine the effects of male and female unemployment rates on fertility in a panel of OECD countries by substituting them individually for overall unemployment rates. The gender-specific unemployment rates and wages used in each of these studies are potentially problematic for a number of reasons. First and foremost is the possibility of endogeneity bias. Gender-specific unemployment rates and wages are likely to be correlated with fertility-induced changes in labor supply, and with unobserved changes in preferences (Hotz, Klerman, and Willis 1997).[9] A second potentially serious problem is that gender-specific unemployment rates are likely to suffer from differing degrees of measurement error. There is evidence that the labor force participation rates of men and women recover at different rates after a recession, signifying differing degrees of "slack" in the economy that are not picked up by measured unemployment rates (Bradbury 2005). Finally, it should be noted that replacing general unemployment rates with either male or female unemployment rates alone prevents identification of the distinct effects of the two variables, as each is likely to serve as a proxy for overall unemployment rates when labor market conditions facing the opposite sex are not controlled for.

As a result of these identification challenges, recent efforts have focused on identifying the income effects of shocks to housing wealth (Lovenheim and Mumford 2013, Dettling and Kearney 2014), changes in male labor market conditions (Lindo 2010, Black et al. 2013), and involuntary job displacement (Del Bono, Weber, and Winter-Ebmer 2012; Ananat, Gassman-Pines, and Gibson-Davis 2013). Though these studies have provided evidence that negative shocks to family income cause decreases in fertility, their methods have not been well suited to identifying the substitution effects from changes in women's labor market opportunities.

In this paper, the indices used to proxy for gender-specific labor market conditions are constructed so as to capture variation in labor demand resulting from exogenous shifts in aggregate product demand that differentially affect men and women based on their likelihood of being employed in the affected industries. The intuition behind this approach is that the degree to which national demand shocks within a specific industry will impact the labor market conditions faced by a given individual depends on the concentration of that industry in the state in which the individual lives and the level of representation of that individual's demographic group (defined primarily as

9. If, for example, a change in preferences were to cause a simultaneous decrease in fertility and an increase in female labor force participation, the fertility rate would decrease and the denominator of the unemployment rate would increase. Holding the number of available jobs fixed, this would cause the coefficient on unemployment to be biased upward (toward zero in the case of a negative relationship).

gender-by-education group) within the industry. Exploiting these plausibly exogenous sources of variation allows me not only to isolate the effects of labor demand shifts but also to identify the effects of changes in women's labor market conditions separately from those of men.

III. Data and Empirical Approach

The fertility data used in this analysis are from the National Center for Health Statistics Division of Vital Statistics (NCHS-DVS). The NCHS-DVS birth certificate data include the near-universe of births occurring in the United States from 1980–2010. I use date of the last menstrual period to determine the date of conception and collapse the data into cells defined by mother's state of residence, and year of conception.[10] Available demographic data include age, race, and education of the mother. I merge the vital statistics natality data with data on state unemployment rates in the year of conception from the Bureau of Labor Statistics (BLS), state employment rates from the Bureau of Economic Analysis (BEA), constructed labor demand indices (described in detail below), and data on state demographic composition by race, age, and educational attainment from the basic monthly Current Population Survey (CPS).

State-level population estimates by race and age come from the National Cancer Institute's Surveillance, Epidemiology and End Results Program (SEER). The SEER data are a modification of vintage 2009 annual population estimates by age, sex, single-race, and Hispanic origin from the U.S. Census Bureau's Population Estimates Program. Because they are based on projections from the decennial census, SEER population estimates are less likely to suffer from sampling bias than population estimates from the CPS. I use the data to construct population denominators for total birthrates as well as birth rates by age and race. Because the SEER data does not include population estimates by level of educational attainment, to construct education-group-specific birth rates, I obtain population estimates by taking the 12-month average of estimated population from the Basic Monthly CPS. I construct birth rates by dividing the number of births by the appropriate at-risk population (females aged 16–45 in the relevant demographic group). Making use of all available data, I have 1,530 state-year observations.

Mean birth rates and unemployment rates for each state over the sample period are presented in Table 1. Distinct regional patterns in fertility are immediately visible. States in the Southwest (Utah, Arizona, New Mexico, Texas, and California) have the highest average birth rates, while the New England states (Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont) have fertility rates among the lowest in the country. Average unemployment rates over the sample period vary widely across states as well, ranging from under 4 percent in Nebraska and South Dakota to over 8 percent in Michigan and West Virginia. The statistics in this table suggest that the demographic makeup of the population and other unobservable state

10. Beginning with the 2005 data, the public-use U.S. microdata natality file no longer includes mother's state of residence. Access to the restricted-use data with geographic identifiers was obtained with special permission from the National Association for Public Health Statistics and Information Systems (NAPHSIS) and the NCHS. Data on the date of last menstrual period is missing for a few state-years. When these data are missing, I impute the date of conception by subtracting nine months from the date of birth.

Table 1
Average Birth and Unemployment Rates by State

State	Birth Rate	Unemployment Rate	State	Birth Rate	Unemployment Rate
Alabama	63.73	6.61	Montana	65.32	5.78
Alaska	79.94	7.88	Nebraska	69.71	3.53
Arizona	78.43	5.89	Nevada	72.70	6.13
Arkansas	67.14	6.46	New Hampshire	57.08	4.30
California	73.73	7.01	New Jersey	62.39	5.83
Colorado	66.27	5.30	New Mexico	75.27	6.75
Connecticut	58.50	4.91	New York	61.72	6.27
Delaware	63.73	4.75	North Carolina	62.89	5.53
District of Columbia	57.15	7.49	North Dakota	66.87	4.10
Florida	64.92	5.92	Ohio	62.62	6.78
Georgia	66.93	5.48	Oklahoma	69.44	5.37
Hawaii	73.37	4.35	Oregon	63.06	7.06
Idaho	77.04	5.90	Pennsylvania	58.61	6.35
Illinois	67.18	6.89	Rhode Island	56.18	5.98
Indiana	64.37	5.95	South Carolina	64.04	6.27
Iowa	64.02	4.82	South Dakota	73.87	3.81
Kansas	70.08	4.77	Tennessee	61.70	6.51
Kentucky	61.97	6.95	Texas	76.70	6.24
Louisiana	69.44	7.30	Utah	94.88	4.97
Maine	55.83	5.52	Vermont	55.09	4.59
Maryland	61.68	5.03	Virginia	61.45	4.47
Massachusetts	56.45	5.32	Washington	64.72	6.84
Michigan	62.12	8.08	West Virginia	56.20	8.65
Minnesota	64.90	4.90	Wisconsin	62.20	5.44
Mississippi	69.84	7.90	Wyoming	69.88	5.23
Missouri	64.70	5.87	United States	66.04	6.19

Sources: Natality data are from the National Center for Health Statistics, Division of Vital Statistics. Population estimates are from the National Cancer Institute's Surveillance, Epidemiology, and End Results Program. Unemployment rates are from the Bureau of Labor Statistics.
Notes: Birth rates are births per thousand women aged 16–45.

characteristics may be important factors in determining both aggregate fertility rates and overall labor market conditions, highlighting the importance of controlling for state fixed effects and nonlinear changes in state demographic composition in the regressions.

Figure 1 shows trends in the birth rate and unemployment rate at the national level over the sample period. The U.S. birth rate appears to follow a procyclical pattern in the national time-series data, reaching a peak of 70.4 births per 1000 women aged 16–45 in 1989, then dropping all the way to 63.3 in 1996 before rebounding all the

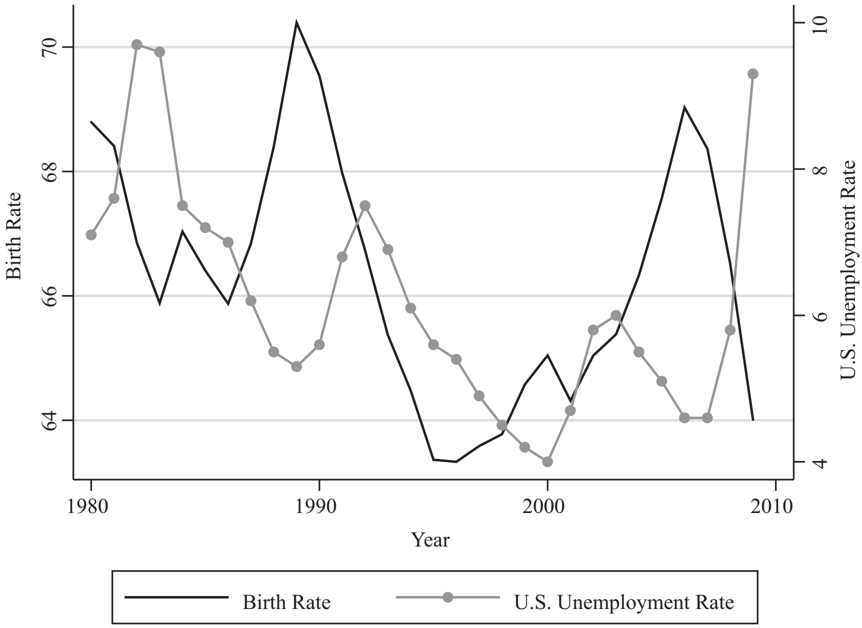

Figure 1
U.S. Birth Rate and Unemployment Rate

Sources: Natality data are from the National Center for Health Statistics, Division of Vital Statistics. Population estimates are from the National Cancer Institute's Surveillance, Epidemiology, and End Results Program. Data on the U.S. unemployment rate are from the Bureau of Labor Statistics.

Notes: Birth rates are calculated as births per thousand women aged 16–45.

way back up to 69.1 in 2006 and dropping sharply again to 64.0 in 2009. Scatter plots of the data, shown in Figure 2, also suggest that fertility is negatively correlated with unemployment. While Panel A shows a very slight positive correlation between current-period unemployment and fertility across state-year observations in the pooled cross-section, there is a marked shift in the relationship when looking at five-year changes in the same variables. Moreover, when focusing on the severe recessionary periods from 1980–82 and from 2007–2009, as in Panels C and D, there is a strong negative correlation between the size of the increase in unemployment in a state over the period and the decline in birth rates over the same period. In fact, though the recessionary period from 2007–2009 displays changes in both unemployment and birth rates that are larger in magnitude than the changes during the recessionary period from 1980–82, the slope of the fitted line is strikingly similar between the two periods.

As the raw time-series and cross-sectional correlations illustrated above are likely to be subject to significant omitted-variables bias, I next turn to regression analysis to estimate the relationship between birth rates and local economic conditions. For the basic regressions, I begin with the following fixed-effects specification:

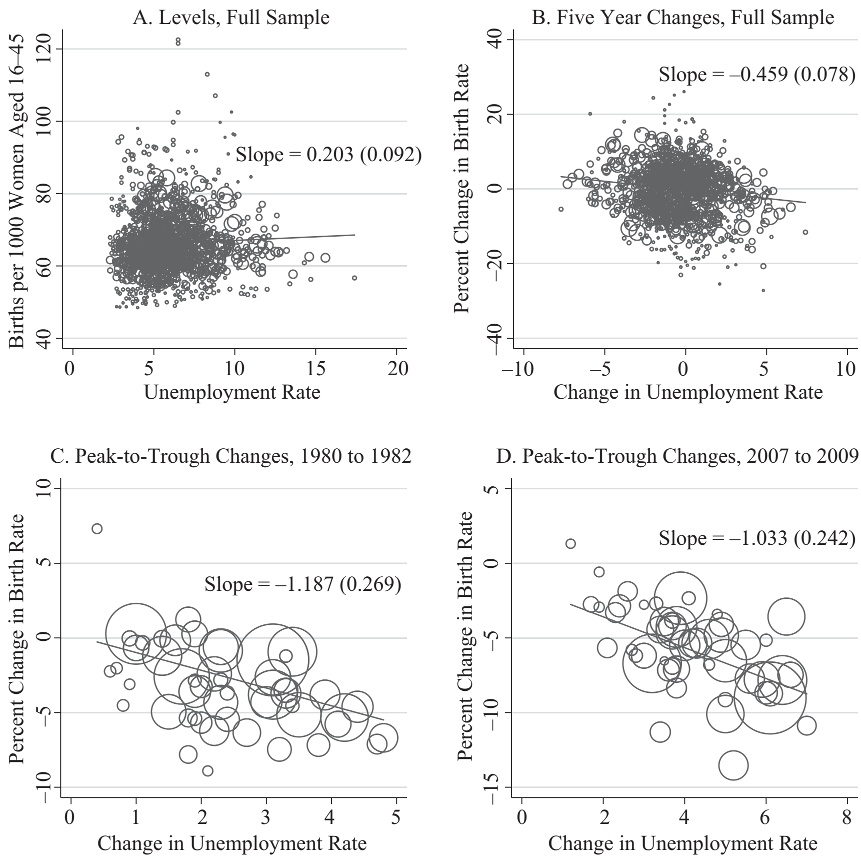

Figure 2
Change in Unemployment Rate versus Percent Change in Birth Rate

Sources: Natality data are from the National Center for Health Statistics, Division of Vital Statistics. Population estimates are from the National Cancer Institute's Surveillance, Epidemiology, and End Results Program. Unemployment rates are from the Bureau of Labor Statistics.

Notes: In Panels A and B, each marker represents a state-year observation. In Panels C and D, each marker represents one U.S. state. Observations are weighted by the female population aged 16–45 in the current or base year. Birth rates are births per thousand women aged 16–45.

$$(1) \quad \ln(Y_{st}) = \beta U_{st} + \psi X_{st} + \alpha_s + \gamma_t + \omega_s * T + \varepsilon_{st}$$

where Y_{st} is the birth rate in state s following year of conception t, and U_{st} is the state unemployment rate in year t. State fixed effects, α_s, are included to control for fixed differences in birth rates across states due to unobservable factors, and year fixed effects, γ_t, are included to account for changes in birth rates over time that are common to all states. State linear time trends, $\omega_s * T$, are included in some specifications to control for unobserved variables correlated with birth rates that change linearly over time within states. Because changes in the demographic composition of a state's population are likely to be correlated both with labor market outcomes and with aggregate fertility rates, my preferred specification also includes time-varying state-level demographic controls, X_{st}, to account for nonlinear changes in population composition by age, race, ethnicity, and educational attainment. Finally, ε_{st} is a random error term. Estimated standard errors are robust and clustered to account for the fact that the error term may be correlated across time periods within each state. The regressions are weighted by the relevant population of women aged 16–45 in each state-year cell.

The use of aggregate unemployment rates as an explanatory variable in this setting has both pros and cons. Though commonly employed as a proxy for local economic conditions and less likely to be endogenous to fertility decisions than individual wages and family income, unemployment rates are problematic in that they capture changes in labor supply as well as changes in labor demand. This increases the likelihood that changes in unemployment will be correlated with changes in other unobserved variables that may also be related to fertility. There also may be a direct reverse-causality bias. If exogenous increases in fertility cause a decline in women's labor force attachment, the denominator of the unemployment rate will decline and, if total employment remains fixed, the measured unemployment rate will increase. As a result, OLS coefficients may be biased upward. Another potential source of bias is measurement error: Unemployment rates are a noisy measure of actual economic conditions. This is especially true in an economic downturn; because "discouraged workers" (workers who want to be employed but are no longer actively searching for a job) are not counted in measured unemployment rates, the unemployment rate may not be capturing the full extent of a recession.

As an alternative to unemployment rates, I capture shocks to labor demand by creating an index of predicted employment growth. The approach is based on the shift-share model used by, among others, Bartik (1991), Katz and Murphy (1992), and Blanchard and Katz (1992). I create a predicted employment growth rate by weighting the national industry-specific employment growth rates by industry shares in each state in a base period and then summing over industries within each state-year as follows:

$$(2) \quad D_{st} = \sum_{i} G_{it} * \frac{E_{is0}}{E_{s0}}$$

where G_{it} is the growth rate of industry i in year t from the March CPS[11] and (E_{is0}) / $(E_{s0})$is the ratio of industry i employment in state s to total employment in state s from the 1980 census.[12] I use this index to instrument for unemployment rates.

An important condition for instrument validity is that the index of predicted employment growth must be correlated with local area unemployment rates. The intuition behind this first-stage correlation is that aggregate demand shocks in a particular

11. Source: IPUMS CPS (King et al. 2010).

12. Source: IPUMS USA (Ruggles et al. 2010). Additional specifications have been run using state industry employment shares from 1970, 1990, and 2000. The results throughout this paper show little sensitivity to this choice of base year.

Table 2
Economic Conditions and Birth Rates

	Ordinary Least Squares			Instrumental Variables		
	1	2	3	4	5	6
Unemployment rate	–0.001	–0.003	–0.008***	–0.016**	–0.022**	–0.022***
	(0.003)	(0.004)	(0.002)	(0.008)	(0.009)	(0.007)
Percent Hispanic			0.020***			0.023***
			(0.005)			(0.005)
Percent Black			–0.001			0.000
			(0.002)			(0.002)
Percent high school graduate			0.003			0.002
			(0.004)			(0.004)
Percent any college			0.001			0.000
			(0.003)			(0.003)
Percent aged 16–25			0.002			0.001
			(0.003)			(0.003)
Percent aged 26–35			0.010***			0.011***
			(0.003)			(0.003)
Percent aged 36–45			–0.006*			–0.007**
			(0.003)			(0.003)
Percent aged 46–55			0.003			0.005*
			(0.003)			(0.003)
Percent aged 66 plus			0.004			0.003
			(0.004)			(0.003)
State time trends	No	Yes	Yes	No	Yes	Yes
First stage statistics						
Shift-share index				–0.698***	–0.647***	–0.636***
				(0.107)	(0.126)	(0.124)
F-statistic				42.00	26.02	25.88
Probability > *F*				0.000	0.000	0.000

Notes: $N = 1{,}530$. The dependent variable is the log of the number of births per 1,000 women aged 16–45. Standard errors (in parentheses) are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by the number of women aged 16–45 in each state-year cell. All regressions include state and year fixed effects.

industry should have larger effects on individuals who live in states in which the affected industry makes up a greater share of total employment. Consistent with this story, the first-stage regression statistics included at the bottom of Table 2 and the scatter plots of the instrument and the unemployment rate in Panels A and B of Appendix Figure A1 show a strong correlation between the instrument and the unemployment rate. Another condition for instrument validity is that the national employment growth rates by industry are uncorrelated with state-level labor supply shocks. As noted by Blanchard and Katz (1992), this will be the case if there is no industry for which employment is concentrated in one particular state. Finally, it is important that there is sufficient cross-sectional variation in base-period industry composition. To ensure that the last two conditions are verified in the data, I use 17 relatively broad industry categories.[13]

IV. Effects of General Labor Market Conditions on Fertility

Results from both the main fixed effects specification and an instrumental variables specification in which predicted employment growth is used as an instrument for the unemployment rate are presented in Table 2. The coefficients on the unemployment rate are negative and significant at the 1 percent level in the ordinary least squares (OLS) specification that includes both state linear time trends and state-level demographic controls (Column 3) and in all three two-stage least squares (2SLS) regressions (Columns 4–6). According to the OLS results, a one percentage-point increase in the unemployment rate is associated with a 0.8 percent decrease in birth rates. The instrumental variables coefficients are larger in magnitude, indicating that a percentage-point increase in unemployment leads to a 1.6 to 2.2 percent decrease in fertility.

As noted previously, first-stage coefficients and *F*-statistics indicate that the index of predicted employment growth is highly correlated with unemployment rates. The fact that the IV coefficients are larger in magnitude is expected, given the direction of the expected reverse-causality bias. Measurement error in unemployment rates could also be causing OLS coefficients to be biased downward in magnitude. The results in Table 2 also make it clear that controlling for nonlinear changes in a state's population composition is important, as coefficients on fraction Hispanic, and on the age composition of the population are significant determinants of state fertility rates, and including them changes the unemployment coefficient in the OLS specification. Thus, my preferred specification is that in Columns 3 and 6, which include demographic controls and state-specific linear time trends.

In Table 3, I present results from three alternative specifications. First, in Panel A, I replace state unemployment rates with state employment to population ratios. Produced by the Bureau of Economic Analysis, these employment rates are based solely on administrative data, making them less susceptible to measurement error than BLS unemployment rates, which are generated from the Current Population Survey. Results using employment rates are similar to those in Table 2. Next, in Panel B, I experiment with using an alternative measure of industry growth in creating my instrument. The

13. Industry categories are as follows: (1) Agriculture, Forestry and Fishing; (2) Mining; (3) Construction; (4) Low Tech Manufacturing (lumber furniture, stone, clay, glass, food, textiles, apparel, and leather); (5) Basic Manufacturing (primary metals, fabricated metals, machinery, electrical equipment, automobile, other transport equipment (excluding aircraft), tobacco, paper, printing, rubber, and miscellaneous manufacturing); (6) High Tech Manufacturing (aircraft, instruments, chemicals, petroleum); (7) Transportation; (8) Telecommunications; (9) Utilities; (10) Wholesale Trade; (11) Retail Trade; (12) Finance, Insurance, and Real Estate; (13) Business and Repair Services; (14) Personal Services; (15) Entertainment and Recreation Services; (16) Professional and Related Services; (17) Public Administration. The division of manufacturing into low-tech, basic, and high-tech categories follows Katz and Murphy (1992).

Table 3
Alternative Specifications

A. Replacing BLS Unemployment Rate with BEA Employment Rate

	Ordinary Least Squares			Instrumental Variables		
	1	2	3	4	5	6
Employment rate	0.001	0.003	0.007***	0.012*	0.023**	0.024***
	(0.002)	(0.003)	(0.002)	(0.007)	(0.010)	(0.007)
First stage F-statistic				32.16	20.53	19.75
Probability > F				0.000	0.000	0.000
State time trends	No	Yes	Yes	No	Yes	Yes
Demographic controls	No	No	Yes	No	No	Yes

B. Instrumental Variables Estimates: Earnings Growth Instrument

	7	8	9	10	11	12
Unemployment rate	0.020**	0.022**	0.020**			
	(0.007)	(0.008)	(0.005)			
Employment rate				0.017**	0.030***	0.027***
				(0.007)	(0.012)	(0.008)
First stage F-statistic	38.19	27.65	28.98	25.45	10.82	12.93
Probability > F	0.000	0.000	0.000	0.000	0.002	0.001
State time trends	No	Yes	Yes	No	Yes	Yes
Demographic controls	No	No	Yes	No	No	Yes

C. Reduced Form Estimates: Predicted Employment and Earnings Growth

	13	14	15	16	17	18
Predicted employment growth	0.011*	0.014**	0.014***			
	(0.006)	(0.006)	(0.004)			
Predicted earnings growth				0.009**	0.009**	0.009***
				(0.004)	(0.004)	(0.003)
State time trends	No	Yes	Yes	No	Yes	Yes
Demographic controls	No	No	Yes	No	No	Yes

Notes: $N = 1{,}530$. The dependent variable is the log of the number of births per 1,000 women aged 16–45. Standard errors (in parentheses) are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by the number of women aged 16–45 in each state-year cell. All regressions include state and year fixed effects. Demographic controls include the fraction of residents who are black, the fraction of residents who are Hispanic, the fraction of residents who are high school graduates, the fraction of residents who are college graduates, and the fraction of residents in each of five age groups: 16–25, 26–35, 36–45, 46–55, and 66 and older.

standard measure of industry growth in shift share analysis is the employment growth rate. However, because changes in industry-specific demand may be reflected not only in employment numbers but also in earnings, in Panel B I try an index that relies on growth in total earnings by industry from the March CPS as the primary measure of industry growth. Results for instrumental variables regressions using this alternate index are similar to the results based on the employment growth indices, though the instrument is slightly weaker, particularly in the employment rate regressions. Finally, in Panel C of Table 3, I include my employment and earnings growth indices directly in the OLS regressions. The results show that both are positively and significantly correlated with birth rates in all three specifications.

Next, I explore differences in effect of labor market conditions on fertility by education, age, and race. As discussed in Section II, there are several reasons why we might expect the responsiveness of fertility to economic shocks to vary across subgroups. However, because there are many different forces at work, it is difficult to predict the nature of these differences. For example, on the one hand, lower-income households are more likely to be credit constrained, so we might expect less-educated, younger, and minority women to reduce their fertility in economic downturns in order to smooth consumption (due to stronger income effects). On the other hand, single women and women with wages below the market rate of childcare may experience stronger substitution effects from wage changes because they are more likely to provide their own childcare. Meanwhile, the predicted effect of differences across groups in the use of contraception and "intentionality" of fertility is ambiguous, as these differences should affect the magnitudes of both income and substitution effects. Another relevant consideration when interpreting differences in the effects across subgroups is that, as shown in recent work by Hoynes, Milles, and Schaller (2012), economic downturns do not affect all workers equally. Less-educated, younger, and minority workers experience larger drops in employment, wages, and work hours in economic downturns for a given increase in the aggregate unemployment rate than other groups. It is important to note, however, that this will not necessarily translate into increased responsiveness of fertility to unemployment shocks for these groups since the same patterns hold for both males and females and thus will affect the magnitude of both the income and substitution effects.

Results from stratified OLS and 2SLS regressions with group-specific fertility rates as the dependent variable are presented in Table 4. The coefficients on the unemployment rate are negative in both the OLS and IV specifications for all subgroups. Though differences in the responsiveness of fertility to economic conditions across groups are not statistically significant, the patterns of the IV coefficients suggest that fertility is in fact less responsive to economic shocks among high school dropouts, younger women, and blacks—the groups that experience the largest cyclical fluctuations in labor market outcomes according to Hoynes, Miller, and Schaller (2012). Thus, it appears that the stronger cyclicality of male labor market conditions facing these groups is counteracted by the stronger cyclicality of female labor market conditions, possibly in combination with larger substitution effects. The patterns in the IV coefficients across subgroups are also not consistent with a story about access to credit, as the least responsive groups are the groups that are the most likely to be credit constrained. One theoretical prediction that is consistent with the pattern of the IV coefficients in Table 4 is that certain groups may experience larger substitution effects from changes in female wage opportunities, either because their wages are below the market rate of childcare or because they are less likely to share childcare responsibilities with a partner. I return to this hypothesis in the next section when I separately examine the responsiveness of fertility to male and female labor market conditions.

Table 4
Effect of Unemployment on Log Birth Rate, by Education, Age and Race

	Ordinary Least Squares			Instrumental Variables		
A. By Education:	Less Than High School	High School Graduate	Any College	Less Than High School	High School Graduate	Any College
Unemployment rate	–0.020***	–0.007	–0.009***	–0.028	–0.048***	–0.037***
	(0.005)	(0.004)	(0.003)	(0.018)	(0.016)	(0.012)
Observations	1,475					
B. By Age:	Age 16–25	Age 26–35	Age 36–45	Age 16–25	Age 26–35	Age 36–45
Unemployment rate	–0.009***	–0.009***	–0.008*	–0.016***	–0.020***	–0.027***
	(0.002)	(0.002)	(0.004)	(0.006)	(0.005)	(0.009)
Observations	1,530					
C. By Race:	White	Black	Other	White	Black	Other
Unemployment rate	–0.008***	–0.008**	–0.011*	–0.023***	–0.012	–0.033*
	(0.002)	(0.003)	(0.006)	(0.008)	(0.009)	(0.017)
Observations	1,530					

Sources: Population estimates by race and age are from the National Cancer Institute's Surveillance, Epidemiology, and End Results Program. Population estimates by education are annual averages from the Basic Monthly Current Population Survey. Notes: Dependent variable is births per 1,000 women aged 16–45 in the relevant demographic group. Standard errors are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by the number of women aged 16–45 in the relevant demographic group in each state-year cell. All regressions include state and year fixed effects, state linear time trends, and demographic controls

V. Effects of Gender-Specific Labor Market Conditions on Fertility

The robust negative correlation between unemployment rates and fertility documented in the previous section is consistent with one major prediction from economic models of fertility: that negative shocks to family income and labor market opportunities are associated with income effects that reduce fertility. However, the procyclical nature of fertility tells us nothing about the relative magnitudes of the income and substitution effects of shocks to female labor market conditions. Contrary to Butz and Ward's (1979) hypothesis and empirical findings for an earlier time period, if declining economic conditions for females do have positive substitution effects, these effects are not large enough to generate countercyclical movements in fertility over the 1980–2009 period. Thus, it remains unclear whether negative shocks to female wages and employment opportunities alone are indeed associated with positive or negative changes in birth rates over the last few decades. It also remains to be seen whether differences in the magnitude of the income and substitution effects across subgroups can provide insight into the observed heterogeneity in the responsiveness of fertility to unemployment shocks. In this section, I address these issues by investigating the response of fertility to demand-induced changes in male and female labor market conditions.

To explore the effects of gender-specific labor market conditions on fertility, I use two separate methods for constructing gender-specific shift-share indices. Both approaches take advantage of a well-documented relationship between changes in national industry employment and sex-specific local labor market conditions (DeBoer and Seeborg 1984).[14] The first builds on the shift-share approach used to create my instrument in the previous section, incorporating changes in aggregate demographic composition within industry in addition to using local industry composition and aggregate shocks to industry-specific employment. Specifically, the indices are created by weighting national industry employment growth rates by the current share of the relevant group within the industry at the national level, as well as the share of that industry in the state in a base period and then summing over industries within each state-year:

$$(3) \quad D_{stg} = \sum_i G_{it} * \frac{E_{igt}}{E_{it}} * \frac{E_{is0}}{E_{s0}}$$

where G_{it} is the national industry employment growth rate, (E_{igt}) / (E_{it}) is the national employment share of gender-education[15] group g in industry i in the current year from

14. See Aizer (2010); Bertrand, Pan, and Kamenica (2013); Blau, Kahn, and Waldfogel (2000); Katz and Murphy (1992); and Qian (2008) for examples of applied research exploiting this relationship.

15. The choice to disaggregate by education group in creating these indices is consistent with previous literature on relative labor demand, which treats different demographic groups (as identified by sex, education, race, and/or experience) as distinct labor inputs that are not perfectly substitutable in the labor market (Katz and Murphy 1992). Analysis by Blank and Gelbach (1991) and Juhn and Kim (1999) suggests that there is little substitution between males and females of the same educational attainment in the labor market. Also, as discussed in Section II, disaggregating by education group allows me to gain insight into differences in the magnitude of the effects of gender-specific labor market conditions on fertility for women that might theoretically be expected to respond differently. However, it should be noted that disaggregating by education group implies an assumption of assortative mating. This assumption is backed by research suggesting that assortative mating is prevalent in the United States. According to Mare (1991), a full 88 percent of marriages

the March CPS, and $(E_{is0}) / (E_{s0})$ is the fraction of state s total employment that is in industry i at the initial period from the 1980 census. For ease of interpretation, I adjust each index so that a one-unit increase is equivalent to a one standard-deviation increase in that index.

The advantage of this approach to constructing gender-specific labor demand indices is that, in addition to exploiting differences in industry growth rates and state employment shares, it uses variation in demographic composition across industries and allows for the gender-education group shares of employment in each industry to change at the national level over time, accounting both for the increase in women's labor force participation and the increase in the size of the college-educated labor force over the sample period. The idea is that, over time, changes in social norms and available technology as well as shifts in product demand alter the gender-education composition of industries. As the shares of women and more-educated workers in particular industries increase over time, individuals in those groups will be more sensitive to national demand shocks in those industries. Meanwhile, these national shocks are dispersed geographically depending on the industry concentration of each state, generating additional variation in labor demand that is exogenous to the fertility and labor supply decisions of individuals in a particular state.

While this first approach accounts for changing demographic composition of industry employment, it has the disadvantage that the magnitudes of the resulting indices are not easily interpreted. As an alternative approach, I create a second set of indices that fix gender-industry shares in the base period and can be more clearly interpreted as "predicted employment growth rates" as in the previous section. These indices are similar to those used in Aizer (2010) and Bertrand, Pan, and Kamenica (2013). Specifically, these indices are created by weighting national industry growth rates, G_{it}, by the fraction of individuals from group g who are working in industry i in state s in 1980, $(E_{igs0}) / (E_{gs0})$, and summing across industries for each state-group-year observation:

$$(4) \quad D_{stg} = \sum_i G_{it} * \frac{E_{igs0}}{E_{gs0}}.$$

These indices represent a predicted employment growth rate for group g in state s in year t. Again, as the only time variation is generated by national-level changes in industry employment growth, the indices are unlikely to be correlated with local area fertility shocks.

Because all of the gender-specific labor demand indices described above exploit the same national industry employment growth over time (weighted differently for males and females based on the representation of each gender within different industries), there is inherent comovement in the male and female shift-share indices. Any joint variation in the two variables (for example, due to business cycle fluctuations) cannot contribute to identification of separate effects of male and female labor demand, so it is important that there is also independent variation in the two variables. This independent variation is illustrated in Panels C and D of Appendix Figure A1, which show scatter plots of both unadjusted and adjusted (residual) values of the predicted employ-

are between spouses who differ by at most one education level. Garfinkel, Glei, and McLanahan (2002) suggest that only 22 percent of married couples have educational attainment that differs by more than two years.

ment growth indices for males and females, overlaid on a 45-degree reference line. The scatter plots also reveal other interesting patterns in the gender-specific demand indices. In particular, predicted male employment growth appears to be weaker than predicted female employment growth for a majority of the state-year observations in the pooled cross-section, as the bulk of the markers falls to the right of the 45-degree line in the unadjusted plot. This pattern is less pronounced when the indices have been adjusted to account for fixed effects, trends, and demographic controls, which suggests that it may reflect the long-term decline of some traditionally male industries over the sample period. Both the unadjusted and residual graphs also show that predicted employment growth for males tends to be weaker than that of females when the economy is experiencing a downturn.[16]

Regression coefficients for the first set of gender-education-specific labor demand indices (the "time-varying group shares" approach) are presented in Panel A of Table 5. Across all groups, the coefficients consistently indicate that an improvement in labor market conditions for men increases fertility rates, while positive labor demand shocks for women are associated with a decline in fertility. For women with less than a high school education, a one standard deviation increase in the labor demand index for men in the same education group is associated with a 4.4 percent increase in current-period fertility, while a one-standard-deviation increase in the corresponding female labor demand index is associated with a 3.1 percent decrease in birth rates. Across education categories, the effects of male labor market conditions remain positive and significant, at 5.6 percent and 3.7 percent for the high school graduate and college-educated groups respectively. Meanwhile, the negative effects of female labor demand shocks decrease both in magnitude and significance across education categories, which suggests that the substitution effects from improved labor market conditions are the strongest for the least-educated women.

Results using the second set of gender-education-specific labor demand indices (the "predicted employment growth" approach) presented in Panel B of Table 5 are consistent with those from the previous table. Again, improvements in male labor market conditions are associated with increased birth rates across all three education groups and improvements in female labor market conditions have negative, insignificant effects. A one percentage point increase in predicted employment growth for men is associated with a 1.7 to 2.3 percent increase in birth rates among women of the same level of educational attainment, while a one percentage point increase in predicted employment growth for women is associated with a 0.2 to 1.1 percent decrease in birth rates (though not statistically distinguishable from zero).

The results in both panels of Table 5 are consistent with the predictions from economic models of fertility, which suggest that increases in family income should increase fertility while increase in female wages may have the opposite effect. Though differences in the coefficients across education groups are not statistically significant, the patterns in the coefficients provide insight into the increasing magnitude of the unemployment coefficients across education groups seen in Table 4. In particular,

16. This finding is consistent with recent work by Hoynes, Miller, and Schaller (2012) and Sahin, Song, and Hobijn (2010), who argue that the dramatic relative decline in labor market conditions for men in the Great Recession is related to differences between males and females in the industry-composition of their employment.

Table 5
Gender-Specific Shift-Share Employment Indices: By Education

	Less Than High School	High School Graduate	Any College
A. Index 1: Time-varying group shares			
Male index, less than high school	0.044**		
	(0.018)		
Female index, less than high school	–0.031**		
	(0.014)		
Male index, high school graduate		0.056***	
		(0.016)	
Female index, high school graduate		–0.021	
		(0.017)	
Male index, any college			0.036**
			(0.014)
Female index, any college			–0.013
			(0.019)
B. Index 2: Predicted employment growth			
Male index, less than high school	0.017***		
	(0.006)		
Female index, less than high school	–0.011		
	(0.008)		
Male index, high school graduate		0.023***	
		(0.006)	
Female index, high school graduate		–0.007	
		(0.007)	
Male index, any college			0.019***
			(0.006)
Female index, any college			–0.002
			(0.013)

Source: Population estimates by education are annual averages from the Basic Monthly Current Population Survey.
Notes: $N = 1{,}475$. Dependent variable is the log of births per 1,000 women aged 16–45 in the relevant education group. Standard errors are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Construction of labor demand indices is described in Section V of the text. All regressions include state and year fixed effects, state time trends and state-year level demographic controls.

they show that the differences across education groups are driven not by increasing responsiveness of birth rates to male labor market conditions as educational attainment increases but rather by greater (negative) responsiveness to female labor market conditions among less-educated women. Returning to the discussion about heterogeneity across subgroups from the previous section, this pattern is consistent with the notion that low-wage women and single women (who make up a larger share of the less-educated groups) experience stronger substitution effects because they are more likely to provide their own childcare rather than rely on market-based childcare options or a spouse or partner.

To further explore heterogeneity in the effects across groups, I separately examine the responses of birth rates to gender-education-specific labor demand for whites and blacks in Table 6. To the extent that white women and black women differ in their rates of single parenthood and their use of external childcare options, we might also expect differences in the income and substitution effects by race. I construct the indices as in Equation 4, where the groups are defined by race as well as by gender and education, and then regress the birth rate for a specific race-education group on gender-specific labor demand indices for the same group. The results do show heterogeneity in the responsiveness of fertility to gender-specific economic conditions for blacks and whites. While improvements in male labor market conditions are associated with similar increases in birth rates both for white and black women, the decline in birth rates resulting from a positive shock to own labor market conditions is much stronger for black women than for white women. Specifically, a one percentage point increase in predicted employment growth for black females is associated with a 2.4–3.6 percent decrease in birth rates among black women, with the effect significant at the 10 percent level for both the less-than-high-school and high-school-graduate groups. For black women in all three education groups, the magnitude of the negative coefficient on the female index is approximately equal to or larger than the positive coefficient on the male index.

The results in Table 6 should be interpreted with caution, as they do not account for unobservable differences in the composition of the groups (for example, by wages, marital status, or age), or for differences by race in assortative mating or the ability to respond to labor market shocks by changing occupation, industry, or geographic location. Moreover, though these results are again consistent with the notion that women with lower wages, higher rates of single parenthood, and less use of market-based childcare, should experience stronger substitution effects, it is also impossible to rule out heterogeneity in preferences across groups. That said, returning to the results from Table 4, the stronger negative response of fertility to changes in own labor market opportunities among black females does provide a potential explanation for the smaller magnitude of the unemployment coefficient for blacks. Further exploration of racial differences in the responsiveness of fertility to changes in labor market opportunities is left to future work.

VI. Robustness

An advantage of the panel data approach used in this paper is that the inclusion of state fixed effects and state-specific time trends accounts for a host of unobservable state characteristics that may be correlated with both fertility and economic conditions. However, these fixed effects control only for variables that are fixed over time or trending linearly within a state. Welfare reform, which occurred in the middle of the sample period under consideration, is a potential source of systematic variation in these unobservable characteristics. With the enactment of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) in 1996,

Table 6
Gender-Specific Predicted Employment Growth and Fertility, by Education and Race

	Less Than High School	High School Graduate	Any College
	A. White		
Male index, white	0.015**		
less than high school	(0.006)		
Female index, white	–0.009		
less than high school	(0.007)		
Male index, white		0.023***	
high school graduate		(0.007)	
Female index, white		–0.006	
high school graduate		(0.010)	
Male index, white			0.017**
any college			(0.006)
Female index, white			0.001
any college			(0.012)
	B. Black		
Male index, black	0.021***		
less than high school	(0.007)		
Female index, black	–0.036**		
less than high school	(0.015)		
Male index, black		0.027***	
high school graduate		(0.008)	
Female index, black		–0.024**	
high school graduate		(0.010)	
Male index, black			0.015
any college			(0.018)
Female index, black			–0.025
any college			(0.023)

Source: Population estimates by race and education are annual averages from the Basic Monthly Current Population Survey.
Notes: $N = 1{,}475$. Dependent variable is the log of births per 1,000 women aged 16–45 in the relevant education group. Standard errors are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Construction of labor demand indices is described in Section V of the text. All regressions include state and year fixed effects, state time trends, and state-year level demographic controls.

states were granted considerable flexibility in setting policies regarding welfare time limits, earnings disregards, sanctions for noncompliance, and the generosity of child-care subsidies—policy parameters that may have important implications for the state economy and the fertility of low income households. To address this issue, I explore the robustness of my results to the inclusion of separate state fixed effects for the pre- and post-welfare reform eras and to allowing the slope of the state linear trends to change in 1996 (when welfare reform occurred at the national level). For the analysis by gender, I focus on individuals with a high school degree or less, as less-educated individuals are most likely to be affected by a changing welfare regime. The results of this analysis, presented in Appendix Table A1, show that the estimated effects of unemployment rates on fertility are robust to these specification changes. Although the coefficient on male labor demand does decrease in magnitude in the final column, the overall pattern of the gender-specific results remains the same.

I have also checked the robustness of the results to changing the definition of the date of conception. I use the date of the last menstrual period to identify the year of conception. However, it is possible that there are systematic differences across groups in the likelihood of accurately reporting this information. As an alternative, I use the date of birth minus nine months. Though there may be nonrandom measurement error in this variable as well due to systematic differences in the average length of gestation across groups, it is reassuring that the results using the two definitions are almost identical. The results using the alternative date of conception are available from the author upon request.

VII. Conclusion

This paper makes two major contributions to the economic literature on fertility. The first contribution is a set of new estimates, based on current best-practices in economic area studies, of the relationship between aggregate economic conditions and birth rates. The second is an empirical test of economic models of fertility, which suggest that improvements in men's labor market conditions should increase fertility while improvements in women's labor market opportunities could potentially decrease birth rates.

To examine movements in fertility over the business cycle, I first estimate the relationship between local labor market conditions and fertility using the standard proxy for local labor market conditions: unemployment rates. My results suggest that fertility is negatively correlated with unemployment rates, with a one percentage point increase in unemployment rates associated with a 0.8 percent decrease in birth rates, or a decrease of 0.53 births per thousand women aged 16–45. Because unemployment rates are likely to be correlated with labor supply as well as labor demand, I also estimate regressions in which I instrument for unemployment rates using a shift-share index of labor demand. As expected, IV estimates of the relationship between unemployment rates and fertility are larger than OLS estimates. According to the IV results, a one percentage point increase in unemployment is associated with a 2.2 percent decrease in birth rates, or 1.46 births per thousand women aged 16–45. The magnitude of these effects is plausible. In recent years, the national annual unemployment rate climbed from a low of 4.6 percent in 2007 to a peak of 9.6 percent in 2010, an increase of five percentage points. The results from this paper suggest that an exogenous increase in unemployment of this size should be associated with a decrease in aggregate birth rates of between 4 percent (OLS) and 11 percent (IV). According to the CDC, the national birth rate for women aged 16–45 actually fell by 7.5 percent between 2007 and 2010 (Martin et al. 2013).

To examine the relative magnitudes of the income and substitution effects of shocks to male and female labor market conditions, I create two different sex-specific indices of labor market demand, one that allows group shares of industry employment to vary over time and one that fixes them at base-period levels. Results using both sets of indices are consistent with the implications of economic models of fertility: Demand-driven improvements in potential wages and/or employment opportunities for men are associated with increased birth rates, while better labor market conditions for women have either no significant effect or a smaller negative effect on birth rates. To the extent that there are negative substitution effects from improvements in female labor market conditions on birth rates, they seem to be concentrated among women with lower levels of educational attainment who face lower wages and are less likely to purchase childcare in the market. Substitution effects also appear to be stronger for black women than for white women.

In this paper, as in many other papers examining the effects of economic conditions on individual outcomes (for example, Ruhm 2000 and the subsequent literature on the health effects of recessions), the mechanisms behind the effects of aggregate economic indicators such as the unemployment rate are difficult to identify. Though much of the effect is likely to operate directly through changes in labor market opportunities (employment, wages, and/or work hours), it is likely that other factors, including changes in household wealth and individual perceptions of the state and stability of the economy may play a role. A mechanism that may be of particular importance in the study of fertility is the price and availability of childcare. It is possible that childcare may become more affordable in an economic downturn, either because of an increase in labor supply to the sector or a decrease in aggregate demand for childcare services. This would work in the opposite direction of the main effect, causing the negative relationship between unemployment rates and birth rates to be smaller in magnitude. Turning to the gender-specific analysis, it is possible that an improvement in female labor demand could result in increasing childcare prices, as increased demand in other sectors might pull labor supply away from the childcare sector while also increasing demand for childcare services. This could partially explain the negative association between female labor demand and fertility. Although the identification strategy used in this paper is not well-suited to exploring what is inside the "black box," the potential for cyclical fluctuations in childcare prices is an interesting avenue for future work.

Whatever the mechanisms behind the effects, the results from the gender-specific analysis suggest that declining economic conditions for females might be expected to result in slightly increased fertility. However, given that the magnitudes of the effects of shocks to male economic conditions are larger in absolute value than the female effect, the procyclical nature of fertility can be explained by transitory fluctuations in men's earnings. This effect should be enhanced if employment and wages in male-dominated industries suffer more in recessions than those in female-dominated industries, which has been the case not only in the "Great Recession" but also in previous recessions (Elsby, Hobign, and Sahin 2010; Hoynes, Miller, and Schaller 2012). An alternative explanation, supported by the findings of Lovenheim and Mumford (2013) and Dettling and Kearney (2014), suggests that the impacts of economic downturns on family nonlabor income, and on housing wealth in particular, may also drive the negative relationship between unemployment rates and fertility.

Appendix

Table A1
Economic Conditions and Birth Rates, Adding Post-96 Interactions

	Ordinary Least Squares			Instrumental Variables		
	1	2	3	4	5	6
	A. Unemployment rates					
Unemployment rate	–0.008***	–0.009***	–0.013***	–0.022***	–0.027***	–0.023**
	(0.002)	(0.002)	(0.001)	(0.007)	(0.008)	(0.010)
State x post 1996	No	Yes	Yes	No	Yes	Yes
State time trend x post 1996	No	No	Yes	No	No	Yes
	Index 1 Time Varying Group Shares			Index 2 Predicted Employment Growth		
	7	8	9	10	11	12
	B. Gender-specific labor demand					
Male index, high school or less	0.045***	0.042**	0.025**	0.019***	0.018***	0.012***
	(0.012)	(0.012)	(0.011)	(0.003)	(0.004)	(0.004)
Female index high school or less	–0.015	–0.011	–0.015	–0.009*	–0.007	–0.010**
	(0.013)	(0.011)	(0.011)	(0.005)	(0.005)	(0.004)
State x post 1996	No	Yes	Yes	No	Yes	Yes
State time trend x post 1996	No	No	Yes	No	No	Yes

Notes: Dependent variable in Panel A is the log of births per 1,000 women aged 16–45. Dependent variable in Panel B is the log of births per 1,000 women aged 16–45 with a high school degree or less. Standard errors are Huber-White robust and clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by the number of women aged 16–45 in the relevant group. All regressions include state and year fixed effects, state time trends, and state-year level demographic controls.

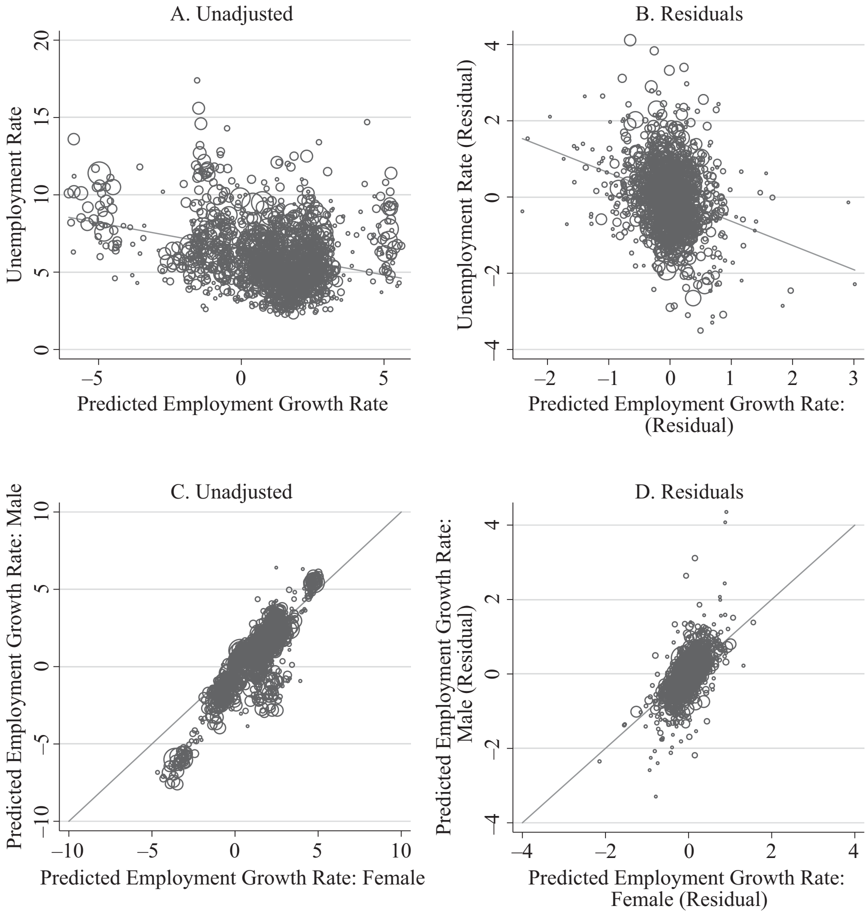

Figure A1

Predicted Employment Growth Instruments

Notes: Each marker represents one state-year observation. Panels B and D show residuals from regressions of the outcome variables on state and year fixed effects, state-specific time trends, and demographic controls. Panels A and B include a fitted regression line. Panels C and D include a 45-degree reference line. Observations are weighted by the female population aged 16–45. Data on state unemployment rates are from the Bureau of Labor Statistics. Construction of the predicted employment growth rate is described in Section IV of the text.

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