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 markets 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-composition within industries, and growth in national industry employment. Consistent with economic theory, I find that improvements in mens labor market conditions are associated with increases in fertility while improvements in womens labor market conditions have the opposite effect. I separately find that increases in unemployment rates are associated with small but significant decreases in birth rates at the state level. JEL: J10, J13, J16.

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1. Introduction

This paper examines the degree to which changes in fertility rates in the United States from 1980 to 2005 can be attributed to demand-induced changes in labor market opportunities for men and women. Economic theory suggests that 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 inspired a large body of economic research examining the response of fertility to exogenous variation in family income and the market-based opportunity cost of women's time. Many economists posit that differences in female labor market opportunities are the source of both the observed cross-sectional relationship between birth rates and income and the long-term decline in fertility that occurred in most developed countries over the 20th century (e.g. Becker (1960), Willis (1973), Ahn and Mira (2002), Jones et al. (2008)). Researchers also cite changes in gender-specific labor demand as an explanation for movements in fertility over the business cycle (see, e.g. Butz and Ward (1979), Macunovich (1995), Orsal and Goldstein (2010)). However, despite the abundance of economic research on the relationship between sex-specific labor market conditions and fertility, there is not consensus about the nature and magnitude of these relationships.

In this paper, I present a new way to separate out the effects of male and female labor market conditions on fertility: I create gender-specific shift-share indices (sometimes referred to as "fixed coefficient" indices) of labor demand that exploit geographic variation in industry concentration, changes in the shares of men and women in each industry over time, and industry employment growth rates at the national level. This is among the first applications of this empirical strategy to the study of fertility in a modern setting.[1] Building on the approach used by Bartik (1991), Katz and Murphy (1992), and Blanchard et al. (1992), this strategy addresses measurement error bias and accounts for the potential endogeneity

[1]Similar approaches have been used to study domestic violence (Aizer (2010)), marriage (Blau et al. (2000)) and crime (Gould et al. (2002)). Schultz (1985) identifies the effect of male and female labor market conditions on fertility using changes in specific world prices and technologies that separately shift male and female labor demand curves in Sweden in the late 19th century. An unpublished paper by Perry (2004) takes a similar approach, producing results that are consistent with those from this paper.

of standard measures of sex-specific labor market conditions. In particular, the shift-share index is more plausibly demand-driven than either gender-specific unemployment rates or observed male and female wages. This analysis by gender provides an important test of the predictions of standard economic models of fertility and sheds light on why fertility moves over the business cycle.

This paper makes a number of other important contributions to the economic literature on fertility. First, to motivate a more careful look at the effects of gender-specific labor market conditions on fertility, I present estimates of the relationship between birth rates and local unemployment rates. These results contribute new estimates to a large literature examining movements in fertility over the business cycle that has yet to come to a consensus.[2] I use a state-year panel and employ a fixed-effects model to control for time-invariant differences in birth rates across states and changes in fertility over time that are common to all states. To account for the potential endogeneity of measured unemployment rates, I also instrument for unemployment rates using a standard industry shift-share labor demand index that exploits geographic variation in industry concentration and national industry employment growth rates. Taking advantage of the demographic detail available in the vital statistics data, I use both OLS and IV approaches to explore differences in the response of fertility to local economic conditions by race, age, educational attainment, and marital status. Finally, I explore the mechanisms behind this main effect with alternate specifications that include of measures of job creation and destruction and aggregate marriage and divorce rates.

Reduced form estimates using gender-specific shift-share indices of labor market demand suggest that improved labor market conditions for men are associated with increases in fertility, while improved labor market conditions for women have the opposite-signed effect. For college graduates, a one standard-deviation increase in the male labor demand index

[2]While most researchers agree that fertility is pro-cyclical (see, among others, Yule (1906), Ogburn and Thomas (1922), Becker (1960), Ben-Porath (1973), Macunovich (1995), Andersson (2000), and Adsera (2005)), there are exceptions. Notably, Butz and Ward (1979) and Mocan (1990), argue that fertility moves countercyclically. Dehejia and Lleras-Muney (2004), find no significant relationship between unemployment rates and birth rates despite observing changes in selection into motherhood over the business cycle.

is associated with a 5.2 percent increase in fertility while a one standard-deviation increase in the female index is associated with 3.7 percent decrease in birth rates. For individuals with less than a high school education, the coefficients are 4.1 percent and -2.8 percent for males and females respectively. For both education groups, the coefficients are statistically distinguishable both from zero and from each other. Meanwhile, results from the general analysis confirm previous empirical findings that increases in local unemployment rates are associated with decreases in birth rates overall. I find that a one-percentage point increase in unemployment rates is associated with a 0.7 to 2.5 percent decrease in birth rates depending on the specification, with the IV specifications yielding coefficients that are larger in magnitude and more precisely estimated. These findings are consistent across educational attainment, racial identification, and marital status, and are robust to, though somewhat attenuated by, the inclusion of measured rates of job creation and destruction and controls for aggregate marriage and divorce rates.

My results are consistent with a key prediction of economic models of fertility: that increases in men's earnings should have a positive effect on fertility. In the presence of uncertainty and imperfect capital markets, the pro-cyclical nature of fertility can be explained by transitory fluctuations in men's earnings. This effect is 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 et al. (2008), Elsby et al. (2010), Hoynes et al. (2012)). Also consistent with economic models of fertility is the result that, holding mens labor-market conditions constant, demand-driven improvements in womens labor market opportunities cause women to substitute away from childbearing. This is among the first papers to provide evidence of this phenomenon using a plausibly exogenous source of demand variation.

This paper will proceed as follows. Section 2 discusses the economic theory and literature related to labor market conditions and fertility. The data and empirical approach are outlined in Section 3. Section 4 presents regression results for both general labor market conditions and gender-specific labor market conditions, and Section 5 contains a discussion of mechanisms. Section 6 concludes.

2. Theory and Identification

Economic theory outlined in early papers by Becker (1960), Mincer (1963), Becker and Lewis (1973), and Willis (1973) has served as the foundation for most empirical analyses of fertility, including those that consider the response of fertility rates to variation in economic conditions. In standard static models of fertility behavior, parents maximize a utility function that depends on quality-adjusted[3] child quantity and all other consumption, subject to a family budget constraint. According to these models, permanent changes in wages, income, and the price of children cause income and substitution effects that alter fertility decisions. In particular, a permanent decrease in male wages is predicted to decrease the total demand for children, while the competing income and substitution effects of a permanent decrease in female wages cause the predicted total effect of the change to be ambiguous.

As impacts of the local labor market fluctuations that I exploit in this paper are likely to be temporary in the long run, more relevant for this analysis are dynamic or life-cycle models of fertility, which assume that individuals can control not only their completed fertility but also the timing of births over their life-cycle (Happel et al. (1984), Hotz et al. (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. Given the identification strategy I use in this paper, I proceed with the dynamic models of fertility in mind. The panel data approach employed

[3]Note that, while the discussion in this paper abstracts from the quantity-quality tradeoff in demand for children, adjustment on the quality margin can be viewed as a substitute for adjustment on the quantity margin and thus would make it more likely that I would find no relationship between local labor market conditions and fertility rates.

here exploits short-term fluctuations in labor market conditions. Furthermore, in the long run individuals will be able to adjust by changing either their industry or their geographic location.

Economic models of fertility do not provide clear predictions about the relationship between fertility and aggregate economic conditions. Movements in unemployment rates are associated with changes in labor market conditions for both men and women, as well as more general changes in the economic environment. Furthermore, to the degree that they are temporary and predictable, business cycle fluctuations will have different effects than unexpected permanent shocks to income or wages. Thus, while estimates of the relationship between birth rates and unemployment rates provide important information about the reduced form contemporaneous effects of business cycles on family decision-making, they should be interpreted with care.

As economic models of fertility choice emphasize important differences between the effects of changes in male and female wages, a way to more thoroughly understand the contemporaneous relationship between fertility and local economic conditions is to separately examine the effects of men's and women's labor market conditions. However, it has proven difficult to cleanly identify the effects of gender-specific labor market conditions on fertility behavior. In an influential paper, Butz and Ward (1979) include both male annual earnings and the female hourly wage in their analysis. However, as Macunovich (1995) points out, their data on female wages are pieced together haphazardly from a number of sources, and their findings are not robust when estimated using data from the March Current Population Survey. Using data from the European Community Household Panel Survey and employing a Cox proportional-hazard model, Adsera (2005) finds that increases in the difference between female and male unemployment rates are associated with acceleration in the timing of births. Using a panel of 22 OECD countries, Orsal and Goldstein (2010) examine the effects of male and female unemployment rates on fertility by substituting them individually for total unemployment rates and find very little difference in coefficients between the two specifications.

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 et al. (1997)). If, for example, a change in preferences were to cause a simultaneous decrease in fertility and a increase in female labor force participation, the fertility rate would decrease and the denominator of the unemployment rate would increase, causing the coefficient on unemployment to be biased upward (toward zero in the case of a negative relationship). 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.

The indices used to proxy gender-specific labor market conditions in this analysis are constructed so as to capture variation in labor demand resulting from exogenous shifts in product demand. This approach relies on two assumptions: that there is imperfect substitution between gender-education groups in the labor market, and that mobility costs limit the degree to which labor market supply and demand adjust across geographic areas in the short term (Blau et al. (2000), Katz and Murphy (1992)). If this is the case, then national growth in demand within a specific industry should differentially impact the labor market conditions faced by an individual depending on the concentration of that particular industry in the state in which the individual lives and the level of representation of that individual's demographic group (defined here as gender-by-education group) within the industry. Thus, by weighting national industry growth rates both by initial-period state industry shares and by national-level demographic shares within each industry, I am able not only to isolate the effects of labor demand shifts but to identify the effects of changes in women's labor market conditions separately from those of men. Given the importance of sex-specific labor market conditions in economic models of fertility, this is an important contribution.

3. Data and Empirical Approach

The fertility data used in this analysis are Vital Statistics natality data from the National Center for Health Statistics (NCHS). The NCHS birth-certificate data include the near-universe of births occurring in the United States from 1980 to 2006. 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.[4] 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, 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 Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). The SEER data are a modification of Vintage 2009 annual population estimates by age, sex, single-race, and Hispanic origin from the US 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 birth rates as well as birth rates by age and race. Because the SEER data does not include population estimates by education and marital status, to construct birth rates by educational attainment and marital status, I obtain those population estimates by taking the twelve-month average of estimated population from the basic monthly CPS. I construct birth rates by dividing the number of births by

[4]Data on the date of last menstrual period is missing for a few state-years. When this data is missing, I impute the date of conception by subtracting nine months from the date of birth.

the appropriate at-risk population (females age 16-45 in the relevant demographic group). Making use of all available data, I have 1326 state-year observations.

Mean birth rates and unemployment rates 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. These regional patterns suggest that immigration may be an important factor in determining fertility rates, highlighting the importance of controlling for nonlinear changes in state demographic composition. Figure 1 shows trends in birth rates and unemployment rates at the national level over the sample period. Birth rates appear to follow a pro-cyclical 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 again.

Differences in patterns in birth rates across demographic groups, shown in Figure 2, emphasize the importance of investigating how the effect of local economic conditions on fertility varies for different groups. Panel (a) shows that birth rates for the 16-25 age group have been slowly declining since 1990, while birth rates for both the 26-35 and 36-45 age groups have increased steadily over the sample period. Meanwhile, a substantial decrease in birth rates for blacks has caused the difference in birth rates between blacks and whites to decline dramatically between 1980 and 2005. Panel (b) shows that, while birth rates for blacks at the beginning of the sample period are approximately 25 percent larger than those of whites, by the end of the sample period, birth rates for blacks and whites are almost identical. As shown in panel (c), the sample period also saw a slight convergence between the birth rates of married and single women, with birth rates of married women decreasing by about eleven percent and the birth rates of single women increasing by 65 percent between 1980 and 2004. Finally, panel (d) presents trends in birth rates by education group. Birth rates for mothers with less than high school increased dramatically over the sample period, while birth rates for high school graduates decreased and birth rates for mothers with any college remained relatively constant. While these differences in fertility patterns across demographic group could result from compositional changes or changes in social norms that are unrelated to economic factors, it could also be that there are inherent differences in the determinants of fertility across demographic groups that cause their responses to fluctuations in local economic conditions to vary.

To motivate the analysis of the effects of gender-specific labor market conditions on fertility, I begin by estimating the relationship between birth rates and general labor market conditions, as measured by state-level unemployment rates. For the basic regressions, I use the following fixed-effects specification:

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

where Y_{st} is the birth rate and U_{st} is the state unemployment rate at the time of conception.[5] 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, which control for unobserved variables correlated with birth rates that change linearly over time within states, are included in some specifications. 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 non-linear changes in population composition by age, race, ethnicity, and educational attainment. The regressions are weighted by the relevant population of women aged 16 to 45 in each state-year cell.[6] Standard errors are robust and clustered at the state level.

[5]See Appendix A, available from the author upon request, for results for a number of alternative regressions specifications, including regressions in which I replace the state unemployment rate with alternative measures of aggregate economic conditions including the state employment-to-population ratio and the male unemployment rate (Table A.1).

[6]Because population weighting may give disproportionate weight to large states, I also run a set of unweighted regressions. Unemployment coefficients from unweighted regressions, shown in appendix table A.1 (available upon request), are both larger and more precisely estimated than those presented in Table 2.

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 (via the denominator) 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. A direct reverse-causality bias is also possible. Holding the number of available jobs fixed, if exogenous increases in fertility cause a decline in women's labor force attachment, the denominator of the unemployment rate will decline and the measured unemployment rate will increase. As a result, OLS coefficients would be biased upward.

Furthermore, because economic theory of fertility focuses on the effects of variation in individual wages and family income, care is needed in interpretation of coefficients on unemployment rates in the context of economic models. As noted above, unemployment rates capture not only the effect of individual job loss, but also changes in the general economic environment, and several economic studies have found that the effects of individual job loss are opposite in sign to the estimated effects of changes in the overall unemployment rate. 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 a prediction of employment growth. The approach is based on the shift-share model used by, among others, Freeman (1980), Bartik (1991), Katz and Murphy (1992), Blanchard et al. (1992), Bound and Holzer (2000), and Gould et al. (2002). Using data from the March Current Population Survey and the 1980 census, 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:

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

where G_{it} is the growth rate of industry i in year t and $\frac{E_{is0}}{E_{s0}}$ is the ratio of industry i employment in state s to total employment in state s from the 1980 census. I use this shift-share index to instrument for unemployment rates. The instrument is valid if the national employment growth rates by industry are uncorrelated with state-level labor supply shocks. As noted by Blanchard et al. (1992), this will be the case if there is no industry for which employment is concentrated in one particular state. It is also important that there is sufficient cross-sectional variation in base-period industry composition. To ensure that these conditions are verified in the data I use 17 relatively broad industry categories.[7]

4. Effects of General Labor Market Conditions on Fertility

4.1. **Results.** Results from both the main fixed effects specification and the instrumental variables specification are presented in Table 2. Like those of Dehejia and Lleras-Muney (2004), the OLS coefficients in columns one through three are negative and statistically insignificant. However, coefficients on unemployment rates are negative and significant at the one-percent level in the ordinary least squares (OLS) specification that includes both state linear time trends and state-level demographic controls (column four) and in all four two-stage least squares (2SLS) regressions (columns five through eight). According to the OLS results, a one percentage-point increase in unemployment rate is associated with a 0.8 percent decrease in birth rates. The instrumental variables coefficients are larger in magnitude,

[7]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).

indicating that a percentage-point increase in unemployment leads to a 1.4 to 2.5 percent decrease in fertility. First-stage F-statistics indicate that my chosen instrument 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.[8] 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 non-linear 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, even in specifications with state-linear time trends included. Based on the results in Table 2, I proceed using the specifications in columns four and eight, which include both demographic controls and state-specific linear time trends.

Changes in these indices from year to year are primarily driven by changes in global product demand and other exogenous factors that influence industry growth rates. The 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, I experiment with an alternative measure of industry growth. In addition to the standard industry employment growth index, I create 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, presented in Table 3, are generally consistent with my main results.

Because fertility behavior differs dramatically by demographic group (see Figure 2), next I explore differences in effect of labor market conditions on fertility by education, race, age, and marital status. Results from stratified OLS and 2SLS regressions with group-specific fertility rates as the dependent variable are shown in Table 4. Because information on mother's education, race, and marital status are missing for some states in some years, these estimates are based on a more limited sample than the main regression results. Each

[8]Note that the OLS coefficients are also significant in the specification in which overall unemployment rates are replaced with male unemployment rates (Table A.2). This suggests a role for reverse-causality or omitted variables bias related to changes in female labor supply in the main specification.

table includes only those state-years for which no more than 10 percent of the observations have missing information on the relevant variable.

The OLS regression results presented in Panel (A) of Table 4 by education group suggest that the effect of changes in unemployment rates on fertility are strongest for the less-than-high-school education group. By contrast, IV results by education, which again are larger in magnitude, suggest that the effect of increased unemployment on fertility is in fact the largest for high school graduates and those with a college education. Note, however, that the difference between coefficients is not statistically significant. Similarly, the IV regressions in Panel (B) of Table 4 suggest that the magnitudes of the effect of unemployment rates on fertility is increasing with age, though again differences in coefficients between age groups are not statistically significant.

As shown in Panel (C), the IV regressions also suggest that birth rates of single women are influenced far more by changes in local economic conditions than birth rates of married women. The IV coefficient indicates that a one percentage point increase in unemployment rates is associated with a 4.8 percent decrease in birth rates for single women. Because marriage and divorce behavior is also impacted by business cycles (Schaller (2012)), this may be due to changes in selection into marriage over the business cycle. Mechanically, this may also be due to the fact that the group of single mothers has a higher concentration of both younger women and women of lower socioeconomic status, both groups that are highly impacted by business cycles. Another possible explanation is that if married individuals are able to insure against shocks to employment and earnings, they will better able to smooth their consumption and thus changes in unemployment rate should have a smaller effect on their fertility decisions.

Finally, because previous research has shown that the fertility responses to changes in unemployment rates differ for black mothers and white mothers (Dehejia and Lleras-Muney (2004)), I stratify by race. Though the two coefficients are not statistically different from one another, the results in Panel (D1) suggest that the effect of an increase in unemployment on the white birth rate is somewhat larger than the effect on the birth rate for blacks. Because differences in coefficients between whites and blacks might be due to difference in the age composition of the two groups, I also run race-specific regressions using total (age-adjusted) fertility rates instead of measured aggregate fertility rates. Total fertility rates (TFRs) are commonly used in demographic research, and are constructed by adding together age-specific fertility rates. The TFR approximates the number of children a woman would bear if she experienced the current age-specific fertility rates throughout her childbearing years.[9] Panel (D2) shows that when accounting for the age distribution of the population, the effects of unemployment on fertility rates are the same for blacks and whites. Interestingly, the results in both panels contrast with the findings of Dehejia and Lleras-Muney (2004), who find that the effect of increase unemployment on birth rates is larger for black women than for white women (though both coefficients are insignificant), and find a small and significant decrease in the percentage of black babies associated with an increase in unemployment.

4.2. **Robustness Checks.** Though aggregate data makes it difficult to identify factors driving changes in individual behavior, a discussion of the possible mechanisms behind the effect of local labor market conditions on fertility is warranted. As documented in a growing literature studying the effects of business cycles on non-labor market outcomes, changes in labor market conditions have a complicated effect on individual behavior.[10] An important question in this and related papers is, to what extent do the effects of changes in labor market conditions on individual non-economic outcomes operate directly through changes in employment status? In an attempt to separate out the effects of job loss from other effects of economic recession, this section explores the sensitivity of this paper's main result to the inclusion of measures of job creation and destruction. Because research has shown that aggregate marriage and divorce rates, which are intrinsically linked to fertility behavior, are independently related to changes in overall economic conditions (Schaller (2012)), I also test the robustness of the main effect to the inclusion of these variables.

[9]Note that the TFR may be an inaccurate estimate of completed fertility in times when cohorts of women are shifting the lifecycle timing of their fertility.

[10]Most of these papers are health-related. See, for example, Ruhm (2000), Miller et al. (2009), Dehejia and Lleras-Muney (2004), Schaller (2012).

Lindo (2010) explores the effect of individual job displacement on married women's fertility. Using an event study approach and panel data, he finds that women are significantly more likely to have children immediately after husband's job displacement but less likely to have children many years later. To reconcile the seeming discrepancy between Lindo's results and those shown here, I use data from the US Census Bureau's new Business Dynamics Statistics data. These publicly available data include tabulations of job creation and destruction by state from the confidential Longitudinal Business Database. The results shown in Table 5 suggest that factors other than job displacement are mediating the effect of economic conditions on fertility. Specifically, controlling for the job destruction rate in the main OLS specification does not significantly change the coefficient on the unemployment rate. However, controlling for the overall unemployment rate, a ten percent increase in the job destruction rate leads to a 1.3 percent increase in the fertility rate. This finding is consistent with Lindo's finding that job displacement in associated with increased fertility in the short run. The fact that the coefficient on the unemployment rate remains unchanged suggests that the effects of business cycle fluctuations extend beyond the direct impacts of job displacement.

Interestingly the job creation coefficient is also positive, though smaller than that for job destruction. This asymmetry in the effects of aggregate employment changes captured in these two variables on fertility is not surprising. While the job creation variable is likely to be picking up the more general effects of economic growth and its effect on individuals is tempered by in-migration, job displacement is a sudden shock to individuals' employment. This theory is supported by the fact that including the job creation rate in the OLS regression causes the magnitude on the coefficient on unemployment to decline. The job creation rate is picking up some of the variation in overall economic conditions that was previously measured only by the unemployment rate.

As Schaller (2012) demonstrates that both marriage and divorce rates are negatively correlated with overall unemployment rates at the state level, I also explore the robustness of my results to the inclusion of aggregate marriage and divorce rates. As marriage and fertility behavior are inextricably linked, it is possible that these variables may play a mediating role in the relationship between labor markets and birth rates. Table 6 shows the results of regressions in which I control for state aggregate marriage and divorce rates. As expected, marriage rates are positively correlated with fertility, while the coefficient on the divorce rate is negative. The coefficient on the unemployment rate is robust to the inclusion of these controls. If anything, the coefficient on unemployment increases with the inclusion of the divorce rate, suggesting that the decrease in divorce rates accompanying an increase in unemployment might be pushing fertility rates in the opposite direction of the main effect.

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

One possible explanation for the robust negative correlation between overall unemployment rates and fertility documented in the previous section is that the disproportionate negative impact to men's wages and employment conditions that occurs in an economic downturn outweighs any positive effects of decreases in the value of women's time. This would be consistent with the finding that employment and wages in male-dominated industries are more negatively impacted by economic recessions than those in female-dominated industries (Elsby et al. (2008), Elsby et al. (2010)). To explore the effects of gender-specific labor market conditions on fertility, I build on the shift-share approach used to create my instrument in the previous section by creating indices of labor demand for males and females that depend on changes in aggregate demographic composition within industry as well as local industry composition and aggregate shocks to industry-specific employment. The index is created at the state level by weighting national industry-specific 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:

$$D_{stf} = \Sigma_i G_{it} * \frac{E_{ift}}{E_{it}} * \frac{E_{is0}}{E_{s0}} \tag{3}$$

G_{it} is the national industry employment or earnings growth rate (discussed below), $\frac{E_{ift}}{E_{it}}$ is the national employment share of group f in industry i in the current year from the March CPS, and $\frac{E_{is0}}{E_{s0}}$ 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.

This approach takes advantage of a well-documented relationship between changes in national industry employment and sex-specific local labor market conditions (DeBoer and Seeborg (1984)).[11] Over time, changes in available technology and shifts in product demand alter the gender-education composition of industries. Here, I exploit variation over time in demographic shares in each industry, as well as changes over time in industry growth and geographic variation in industry concentration, to identify the effects of gender-specific labor market conditions by education group, and assume that employer preferences, skill-specific technology, and mobility costs limit mobility between industries and geographic areas.

I disaggregate by education group in these sex-specific regressions for a number of reasons. First, if I were to focus only on overall industry shares by gender, the resulting labor demand indices for male and females would be inversely related by definition, forcing a restriction on their coefficients in my regressions. This is not the case when demographic groups are defined on more than one dimension.[12] More importantly, this approach 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 (e.g.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. However, it should be noted that disaggregating by education group implies an assumption

[11]see, e.g. Aizer (2010), Blau et al. (2000), Katz and Murphy (1992), Qian (2008) for examples of applied research exploiting this relationship.

[12]With only two groups, an increase in the share of one group is, by construction, associated with a decrease in the share of the other group within that industry. However, an increase in the share of women with less than a high school education in a particular industry is not necessarily associated with a decrease in the share of men with less than a high school education in that same industry, allowing the two indices to move independently of one another.

of assortative mating[13]. To the extent that women with low educational attainment are matching with men with higher levels of education, or vice-versa, these estimates will not be capturing the full effect of changes in labor market conditions faced by a woman's partner (or potential partners) on her fertility.

Results using these gender-specific labor demand indices on the whole sample are presented in Tables 7 and 8. Table 7 presents coefficients on indices that exploit national industry employment growth rates from the March CPS. Across all groups, the coefficients consistently indicate that an improvement in labor market conditions for men increases fertility rates, while an increase in labor market conditions for women is associated with a decline in fertility, with most coefficients significant at least at the five-percent level. With all indices adjusted so that a one-unit increase is equivalent to a one standard-deviation increase, it appears that the effects of changes in the male labor demand indices are consistently stronger than the effects of changes in female labor demand indices. Differences between coefficients on male and female labor demand indices are statistically significant for all three education groups. Overall, the effects of changes in gender-specific labor market conditions are the strongest for the college-educated group: a one standard-deviation increase in labor demand index for college-educated males is associated with a 5.2 percent increase in the fertility of college-educated women, while a one standard-deviation increase in labor demand for college-educated females is associated with a 3.7 percent decline in their fertility. Evaluated at the mean birth rate for college educated women, the male effect is equivalent to an increase of 2.87 births per thousand college-educated women aged 16-45. The female effect is equivalent to a decrease of 2.04 births per thousand. The same effects for the less-than-high-school group are 4.1 percent (2.72 births) and -2.8 percent (1.86 births) respectively.

As with the general labor demand indices, changes in these sex-specific indices from year to year are primarily driven by changes in global product demand and other exogenous

[13]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 are between spouses who differ by at most one education level. Garfinkel et al. (2002) suggest that only 22 percent of married couples have educational attainment that differs by more than two years.

factors that influence industry growth rates. Because changes in industry-specific demand may be reflected not only in employment numbers, but also in earnings, I experiment with an alternative measure of industry growth. Results from regressions in which industry employment growth is replaced with industry earnings growth rates from the March CPS are provided in Table 8. While the coefficients on the wage growth indices are mostly consistent in sign and magnitude with those on the employment growth indices, only the coefficients for males in the two higher education groups are statistically significant, and I cannot reject equality between the coefficients for males and females.

6. Conclusion

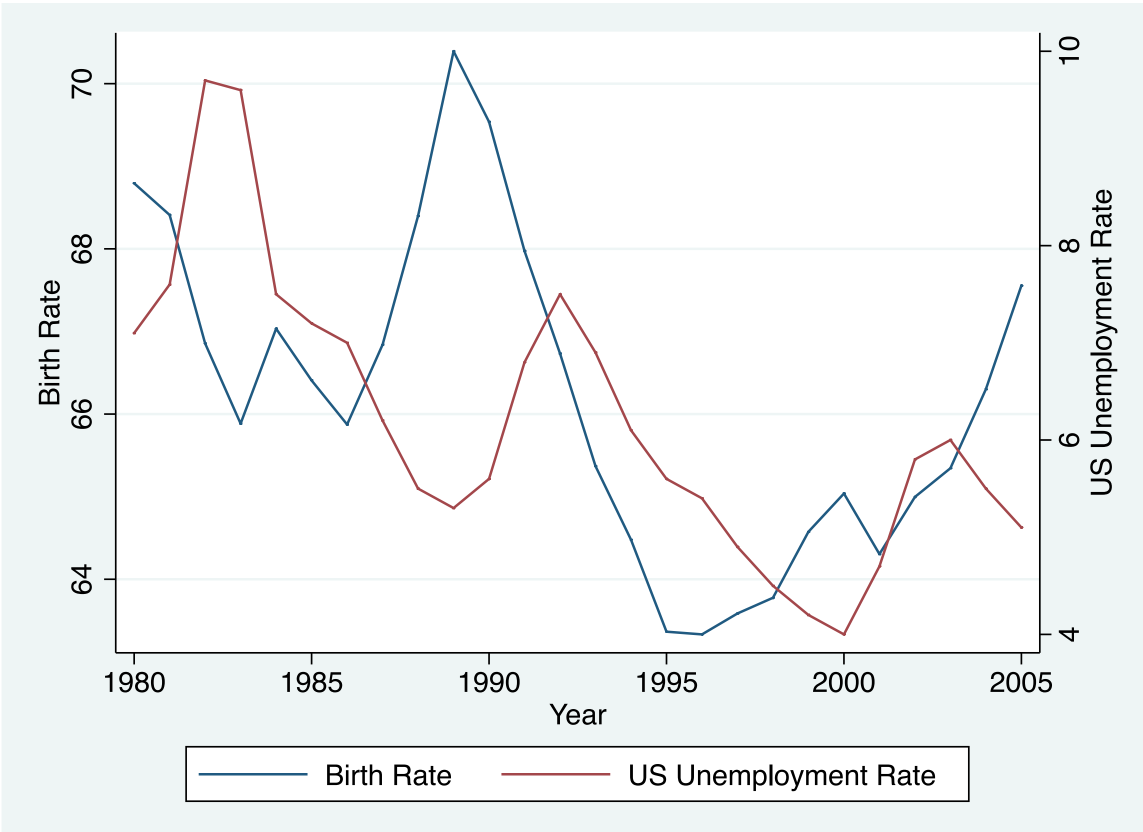
In this paper, I have presented new evidence that birth rates move pro-cyclically. First, I estimate the relationship between local labor market conditions and fertility using the standard proxy for local labor market conditions: unemployment rates. Using a state-year panel of vital statistics data and a fixed-effects specification, I am able to provide estimates that are cleaner, more reliable, and more recent, than were previously available. 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 .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 2.6 percent decrease in birth rates, or 1.71 births per thousand women aged 16-45.

This paper also contains 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.

I test this prediction by creating sex-specific indices of labor market demand, and my results 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 the opposite effect.

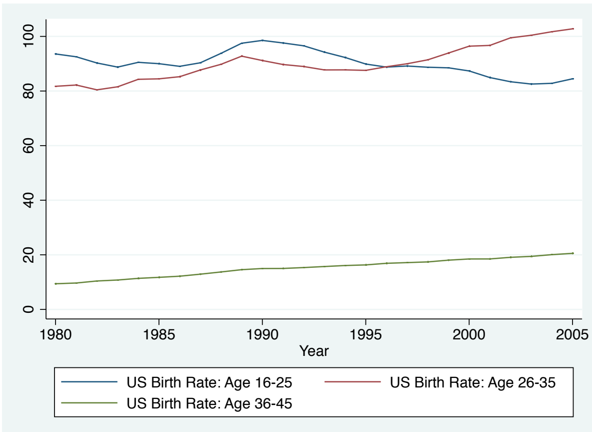
The results from this analysis have important implications for the literature on the cyclicality of fertility behavior. Though many researchers have attempted to isolate the effect of men's and women's labor market conditions on fertility, this paper is the first to create a plausibly exogenous proxy for sex-specific labor market conditions to obtain estimates of the income and substitution effects predicted by economic theories of fertility. My finding of a positive association between male labor market conditions and fertility rates is consistent with a finding of an overall negative association between aggregate unemployment rates and birth rates, as the employment and wages of men have been hit harder in recent aggregate recessions than those of women (DeBoer and Seeborg (1984), Elsby et al. (2010)).

Figure 1. US Birth Rates and Unemployment Rates

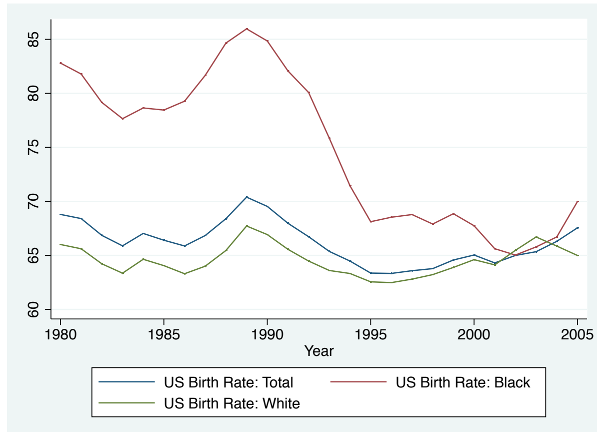


Note: Birth rates are calculated as births per thousand females aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). Data on the US unemployment rate are from the Bureau of Labor Statistics.

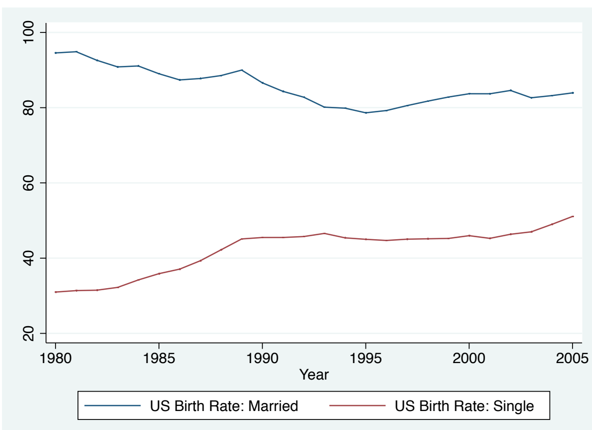
Figure 2. US Birth Rates by Demographic Group



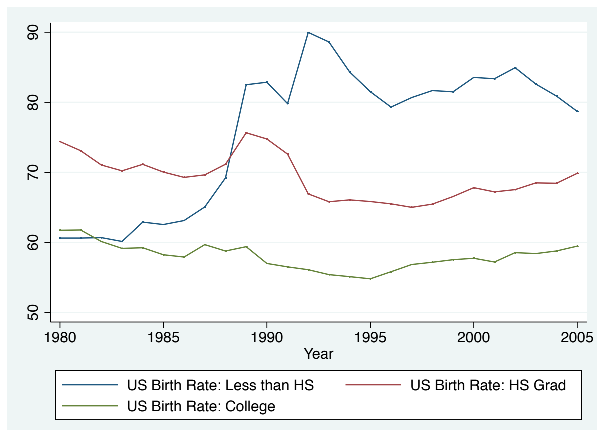
(a) By Age



(b) By Race



(c) By Marital Status



(d) By Education

Note: Birth rates are births per thousand females in the relevant demographic groups aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates by race and age are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). To maintain comparability over time, each graph excludes states for which the relevant variable is missing in any sample year

TABLE 1. **Average Birth Rates and Unemployment Rates by State**

State	Birth Rate	Unemp. Rate	State	Birth Rate	Unemp. Rate
Alabama	63.42	6.77	Montana	65.05	5.99
Alaska	80.22	8.05	Nebraska	68.84	3.55
Arizona	78.90	5.90	Nevada	72.56	6.02
Arkansas	66.70	6.57	New Hampshire	58.03	4.30
California	74.22	6.98	New Jersey	62.12	5.82
Colorado	66.07	5.31	New Mexico	75.40	7.04
Connecticut	58.76	4.78	New York	61.83	6.36
Delaware	63.39	4.71	North Carolina	62.25	5.38
DC	56.60	7.57	North Dakota	66.43	4.20
Florida	65.08	5.90	Ohio	62.55	6.75
Georgia	66.64	5.36	Oklahoma	68.67	5.50
Hawaii	72.98	4.40	Oregon	62.94	7.07
Idaho	76.56	6.08	Pennsylvania	58.52	6.47
Illinois	67.45	6.95	Rhode Island	56.54	5.78
Indiana	63.92	5.88	South Carolina	63.59	6.06
Iowa	63.35	4.87	South Dakota	73.29	3.86
Kansas	69.43	4.75	Tennessee	61.17	6.47
Kentucky	61.38	6.92	Texas	76.49	6.36
Louisiana	69.36	7.69	Utah	95.39	5.11
Maine	56.17	5.50	Vermont	55.75	4.57
Maryland	61.40	5.08	Virginia	61.02	4.52
Massachusetts	56.66	5.26	Washington	64.50	6.97
Michigan	62.55	7.94	West Virginia	55.49	9.17
Minnesota	64.55	4.80	Wisconsin	62.10	5.40
Mississippi	69.59	7.99	Wyoming	68.96	5.43
Missouri	64.47	5.80			

Note: Birth rates are births per thousand females aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). Data on state unemployment rates are from the Bureau of Labor Statistics.

Table 2. Results for Different Regression Specifications

	(1) OLS 1	(2) OLS 2	(3) OLS 3	(4) OLS 4	(5) IV 1	(6) IV 2	(7) IV 3	(8) IV 4
Unemployment Rate	0.000272	-0.00219	-0.00401	-0.00816***	-0.0138*	-0.0241***	-0.0157**	-0.0252***
	(0.0030)	(0.0047)	(0.0031)	(0.0025)	(0.0082)	(0.0093)	(0.0070)	(0.0067)
Pct Hispanic			0.0113***	0.0218***			0.0134***	0.0266***
			(0.0024)	(0.0046)			(0.0028)	(0.0057)
Pct Black			0.00197	-0.00214			0.00322	-0.000608
			(0.0022)	(0.0021)			(0.0022)	(0.0018)
Pct HS Grad			0.00692***	0.00373			0.00607**	0.00403
			(0.0025)	(0.0042)			(0.0028)	(0.0046)
Pct Any College			0.0104***	0.00348			0.00970***	0.00281
			(0.0029)	(0.0032)			(0.0028)	(0.0033)
Pct 16-25			0.00567*	0.00621**			0.00427	0.00504*
			(0.0032)	(0.0027)			(0.0030)	(0.0029)
Pct 26-35			0.0116***	0.00900***			0.0136***	0.0129***
			(0.0037)	(0.0030)			(0.0041)	(0.0038)
Pct 36-45			-0.00301	-0.00302			-0.00387	-0.00274
			(0.0042)	(0.0033)			(0.0044)	(0.0035)
Pct 46-55			0.00520	0.00655**			0.00684	0.0100***
			(0.0040)	(0.0030)			(0.0044)	(0.0037)
Pct 66 plus			0.00845*	0.00699*			0.00750	0.00663*
			(0.0049)	(0.0039)			(0.0046)	(0.0038)
State Time Trends	No	Yes	No	Yes	No	Yes	No	Yes
Observations	1326	1326	1326	1326	1326	1326	1326	1326

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

All regressions include state and year fixed effects.

Table 3. Instrumental Variables Estimates using Earnings Growth Index

	(1) IV 1	(2) IV 2	(3) IV 3	(4) IV 4
Unemployment	-0.0171**	-0.0238***	-0.0169**	-0.0234***
Rate	(0.0082)	(0.0082)	(0.0069)	(0.0056)
F-Stat (1st Stage)	32.37	22.04	26.27	23.14
Prob>F	0.000	0.000	0.000	0.000
State Time Trends	No	Yes	No	Yes
Demographic Controls	No	No	Yes	Yes
Observations	1326	1326	1326	1326

Dependent variable is log of births per 1000 women aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). Data on the US unemployment rate are from the Bureau of Labor Statistics. Standard errors are Huber-White robust and clustered at the state level. * $p < .1$, ** $p < .05$, *** $p < .01$. Estimates are weighted by the number of women aged 16-45 in each state-year cell. All regressions include state and year fixed effects.

Table 4. Birth Rate Regressions by Demographic Group

(A) Regressions by Education Group

	(1) OLS:Ed<12	(2) OLS:Ed=12	(3) OLS:Ed>12	(4) IV:Ed<12	(5) IV:Ed=12	(6) IV:Ed>12
Unemployment Rate	-0.0165*** (0.0054)	-0.0120*** (0.0032)	-0.0108*** (0.0037)	-0.0279** (0.013)	-0.0393*** (0.0090)	-0.0317*** (0.0088)
Observations	1216	1216	1216	1216	1216	1216

(B) Regressions by Age Group

	OLS:1625	OLS:2635	OLS:3645	IV:1625	IV:2635	IV:3645
Unemployment Rate	-0.00816** (0.0033)	-0.00999*** (0.0021)	-0.00545 (0.0034)	-0.0205*** (0.0064)	-0.0213*** (0.0046)	-0.0237** (0.0093)
Observations	1326	1326	1326	1326	1326	1326

(C) Regressions by Marital Status

	(1) OLS:married	(2) OLS:single	(3) IV:married	(4) IV:single
Unemployment Rate	-0.0129*** (0.0022)	-0.0153** (0.0067)	-0.0288*** (0.010)	-0.0476*** (0.013)
Observations	1325	1325	1325	1325

(D1) Regressions by Race

	(1) OLS:white	(2) OLS:black	(3) OLS:other	(4) IV:white	(5) IV:black	(6) IV:other
Unemployment Rate	-0.00804*** (0.0027)	-0.00794** (0.0031)	-0.00904 (0.0058)	-0.0196*** (0.0056)	-0.0154** (0.0073)	-0.0325** (0.016)
Observations	1283	1283	1283	1283	1283	1283

(D2) Regressions by Race: Total (Age-Adjusted) Fertility Rate

	(1) OLS:white	(2) OLS:black	(3) OLS:other	(4) IV:white	(5) IV:black	(6) IV:other
Unemployment Rate	-0.00815*** (0.0020)	-0.00832*** (0.0029)	-0.00941 (0.0056)	-0.0114** (0.0048)	-0.0123* (0.0073)	-0.0290* (0.015)
Observations	1283	1283	1283	1283	1283	1283

Dependent variable is births per 1000 women aged 16-45 in the relevant demographic group. Natality data are from the National Center for Health Statistics. Population estimates by race and age are from the NHIS-SEER. Population estimates by education and marital status are from the CPS. Data on the US unemployment rate are from the Bureau of Labor Statistics. Standard errors are Huber-White robust and clustered at the state level. * $p < .1$, ** $p < .05$, *** $p < .01$. 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.

Table 5. Controlling for Job Creation and Destruction

	(1) Log Birth Rate	(2) Log Birth Rate	(3) Log Birth Rate
Unemployment Rate	-0.00963***	-0.00547**	-0.00653***
	(0.0024)	(0.0025)	(0.0022)
Job Destruction Rate (ln)	0.129***		0.132***
	(0.017)		(0.017)
Job Creation Rate (ln)		0.0582*	0.0678**
		(0.030)	(0.029)
Observations	1326	1326	1326

Dependent variable is log births per 1000 women aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). Data on the US unemployment rate are from the Bureau of Labor Statistics. Data on job creation and destruction are from the Census Bureau's Business Dynamics Statistics. Standard errors are Huber-White robust and clustered at the state level. * $p < .1$, ** $p < .05$, *** $p < .01$. Estimates are weighted by the number of women aged 16-45 in each state-year cell. All regressions include state and year fixed effects, state linear time trends and state-year level demographic controls.

Table 6. Controlling for Marriage and Divorce Rates

	(1) Log Birth Rate	(2) Log Birth Rate	(3) Log Birth Rate
Unemployment Rate	-0.00773***	-0.0107***	-0.00683***
	(0.0024)	(0.0024)	(0.0022)
Log Marriages per 1000 Single Females	0.0302*		0.0273*
	(0.017)		(0.015)
Log Divorces per 1000 Married Females		-0.0386***	-0.0448***
		(0.014)	(0.014)
Observations	1271	1219	1219

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

All regressions include state and year fixed effects, state time trends, and demographic controls.

Table 7. Gender-specific Shift-Share Index: By Education

	(1) lths	(2) hsgrad	(3) college
(sum) index_male_lths	0.0447** (0.017)		
(sum) index_female_lths	-0.0325*** (0.012)		
(sum) index_male_hsgrad		0.0394*** (0.011)	
(sum) index_female_hsgrad		-0.00591 (0.016)	
(sum) index_male_college			0.0510*** (0.015)
(sum) index_female_college			-0.0366** (0.017)
Observations	1164	1164	1164

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

All regressions include state and year fixed effects and state time trends.

Table 8. Gender-specific Shift-Share Index: By Education, CPS Wage Growth Index

	(1) lths	(2) lths	(3) hsgrad
(sum) index_male_lths_wg	0.0284*		
	(0.016)		
(sum) index_female_lths_wg	-0.0219		
	(0.020)		
(sum) index_male_hsgrad_wg		0.0285**	
		(0.012)	
(sum) index_female_hsgrad_wg		0.0179	
		(0.021)	
(sum) index_male_college_wg			0.0369**
			(0.016)
(sum) index_female_college_wg			-0.0155
			(0.037)
Observations	1164	1164	1164

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

All regressions include state and year fixed effects and state time trends.

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Appendix A. Alternative Regression Specifications

Table A.1

	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) IV	(6) IV	(7) IV	(8) IV
Replacing Unemployment Rate with Employment-Population Ratio								
E-Pop Ratio	-0.00135	0.00149	0.00132	0.00577*	0.0145	0.0268**	0.0153**	0.0270***
	(0.0020)	(0.0035)	(0.0027)	(0.0029)	(0.0099)	(0.012)	(0.0071)	(0.0076)
Replacing Overall Unemployment Rate with Male Unemployment Rate								
Male Unemployment Rate	0.00132	-0.000962	-0.00255	-0.00572**	-0.0120*	-0.0207**	-0.0134**	-0.0214***
	(0.0027)	(0.0037)	(0.0028)	(0.0023)	(0.0072)	(0.0081)	(0.0061)	(0.0056)
Unweighted Regressions								
Unemployment Rate	-0.00892***	-0.0114***	-0.0116***	-0.0132***	-0.0240**	-0.0399***	-0.0258***	-0.0345***
	(0.0032)	(0.0031)	(0.0026)	(0.0026)	(0.0099)	(0.010)	(0.0097)	(0.0080)
Replacing Aggregate Fertility with Total (Age-Adjusted) Fertility								
Unemployment Rate	-0.00321	-0.00432	-0.00518*	-0.00896***	-0.00712	-0.0129*	-0.00880	-0.0156***
	(0.0024)	(0.0043)	(0.0028)	(0.0018)	(0.0066)	(0.0073)	(0.0061)	(0.0049)
State Time Trends	No	Yes	No	Yes	No	Yes	No	Yes
Demographic Controls	No	No	Yes	Yes	No	No	Yes	Yes
Observations	1326	1326	1326	1326	1326	1326	1326	1326

Dependent variable is log of births per 1000 women aged 16-45. Natality data are from the National Center for Health Statistics. Population estimates are from the National Health Interview Survey (NHIS) Survey Epidemiology and End Results (SEER). Employment-population ratios are from the Bureau of Labor Statistics. Male unemployment rates are annual averages from the Basic Monthly Current Population Survey. Standard errors (in parentheses) are Huber-White robust and clustered at the state level. * $p < .1$, ** $p < .05$, *** $p < .01$. Estimates are weighted by the number of women aged 16 to 45 in each state-year cell. All regressions include state and year fixed effects.

TABLE A.2. First-Difference Specification

	(1) OLS 1	(2) OLS 2	(3) Inst = index	(4) Inst = index	(5) Inst = indchg	(6) Inst = indchg
D.Unemployment Rate	-0.00817***	-0.00856***	-0.00827***	-0.0103***	-0.00846***	-0.00807***
	(0.00046)	(0.00065)	(0.00038)	(0.00076)	(0.00088)	(0.00097)
L.Pct Hispanic		-0.0000345		-0.0000503		-0.0000302
		(0.00013)		(0.00013)		(0.00013)
L.Pct Black		0.0000452		0.0000368		0.0000475
		(0.000091)		(0.000091)		(0.000090)
L.Pct HS Grad		-0.000573**		-0.000641**		-0.000554**
		(0.00027)		(0.00027)		(0.00027)
L.Pct Any College		0.000394		0.000367		0.000401*
		(0.00024)		(0.00024)		(0.00024)
L.Pct 16-25		-0.00476***		-0.00445***		-0.00484***
		(0.0011)		(0.0011)		(0.0011)
L.Pct 26-35		-0.00551***		-0.00535***		-0.00556***
		(0.0013)		(0.0012)		(0.0013)
L.Pct 36-45		-0.00475***		-0.00459***		-0.00480***
		(0.0011)		(0.0011)		(0.0012)
L.Pct 46-55		-0.00334*		-0.00301*		-0.00343*
		(0.0019)		(0.0018)		(0.0019)
L.Pct 66 plus		-0.00571***		-0.00544***		-0.00578***
		(0.0014)		(0.0014)		(0.0015)
Constant	-0.00156***	0.431***	-0.00156***	0.414***	-0.00158***	0.436***
	(0.00043)	(0.11)	(0.00042)	(0.10)	(0.00043)	(0.11)
Observations	1225	1225	1225	1225	1225	1225

Standard errors in parentheses

* $p < .1$, ** $p < .05$, *** $p < .01$

All regressions include state and year fixed effects.