MALE EARNINGS, MARRIAGEABLE MEN, AND NONMARITAL FERTILITY: EVIDENCE FROM THE FRACKING BOOM

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Abstract—We investigate whether an increase in the potential earnings of men leads to an increase in marriage and a reduction in nonmarital births by exploiting the positive economic shock associated with fracking in the 2000s. A reduced-form analysis reveals that in response to local-area fracking production, which increased wages and jobs for non-college-educated men, both marital and nonmarital birth rates increase, but marriage rates do not. The pattern of results is consistent with positive income effects on births but no associated increase in marriage. We contrast our findings to the Appalachian coal boom experience of the 1970s and 1980s.

I. Introduction

THERE is widespread interest among social scientists and policy observers in the declining rates of marriage among less educated individuals and the concomitant rise in nonmarital childbearing. In 2014, over 40% of all births in the United States were to an unmarried mother, with an even higher rate of 62% among non-college-educated mothers.[1] A leading conjecture as to why so many less educated women are choosing motherhood without marriage points to the weak economic prospects of their male partners. The idea is that changing labor market structures and economic conditions have adversely affected the economic prospects of less educated men, making them "less marriageable." This concept was first posited in the seminal work of Wilson and Neckerman (1986) and Wilson (1987), who were writing about the rise in nonmarital childbearing among African American women in urban settings. The rise in nonmarital childbearing is now much more widespread, reflecting a dramatic, steady shift over the past fifty years, and it is no longer a distinctly urban or minority experience.

Though conceptually compelling and generally consistent with empirical patterns, there is scant empirical evidence about the causal link between male economic status and rates of marriage or nonmarital childbearing. It is very difficult to empirically isolate the causal relationship of interest because high-earning men might be more likely to marry for other reasons, and areas with better male economic opportunities might also be where more marriage- or family-inclined individuals choose to live for other reasons. A recent working paper by Autor, Dorn, and Hansen (2017) provides a notable exception. Exploiting trade shocks during the period 1990 to 2010, the authors find that a decline in male employment opportunities driven by import shocks leads to a decline in births, a decline in marriage, a rise in births to teen mothers, and an increase in the number of children being raised in single-mother families.[2] We investigate a reverse marriageable men hypothesis, asking if an improvement in the economic position of men leads to a positive effect on marriage and a corresponding decrease in nonmarital birthrates.

We attempt to shed light on the causal relationship between improved labor market opportunities for less educated men and subsequent fertility and marriage outcomes by using the fracking boom as a source of improved wage prospects for less educated men. Our empirical analysis is based on data from fracking areas around the country during the period 1997 to 2012, with the deliberate exclusion of North Dakota and Montana.[3] We begin by documenting that localized fracking booms had a sizable effect on the earning potential of less educated males. We then conduct a reduced-form analysis documenting the relationship between local area simulated new fracking production and birth and marriage outcomes. The results of the analysis do not indicate a shift toward marriage in response to an increase in the potential wages of less educated men associated with localized fracking booms. But both marital and nonmarital births increase significantly, consistent with the notion that children are "normal goods," as proposed by Becker (1960), and confirmed in subsequent empirical work (Black et al., 2013; Dettling & Kearney, 2014; Lindo, 2010; Lovenheim & Mumford, 2013). We confirm that the results

Received for publication May 10, 2017. Revision accepted for publication October 23, 2017. Editor: Amitabh Chandra.

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We are grateful to Phil Levine, Kevin Lang, Seth Sanders, Na'ama Shenhav, and Jim Ziliak for helpful comments. We also thank our University of Maryland colleagues for many helpful comments during the applied microeconomics and population center workshops, as well as seminar participants at the University of Virginia, UC Davis, Boston University, Georgetown University, and University of Kentucky. We gratefully acknowledge financial support from a University of Maryland Population Research Center seed grant.

A supplemental appendix is available online at http://www.mitpressjournals.org/doi/suppl/10.1162/rest_a_00739.

[1] For a recent review of the relevant literature and a conceptual framework for thinking about nonmarital childbearing from the perspective of children's outcomes, see Kearney and Levine (2017).

[2] Earlier work offered some corroborating support for the notion. For example, work by Charles and Luoh (2010) finds that increased rates of male incarceration due to policy shifts led to decreased rates of marriage in affected marriage markets. McLanahan and Watson (2011) document that for a given earnings level, men who are of relatively lower income than their peers are less likely to be married. Kearney and Levine (2014) find that conditional on becoming pregnant, low-socioeconomic-status young women are less likely to marry before having the baby if they live in a place with a greater level of income inequality, consistent with a story of their being less likely to find their (presumably low-SES) male partners to be desirable marriage partners. Cherlin, Ribar, and Yasutake (2016) finds that part of that documented relationship appears to be due to fewer middle-skill male jobs in more unequal places. The ethnographic work of Edin and Kafalas (2011) suggests that among their sample of interviewed single mothers, an important factor in their decision not to marry the child's father relates to a perceived lack of economic security that he would bring to the family.

[3] Given the unique experience of fracking in the Bakken region, which was characterized by a large in-migration of male workers into sparsely populated communities, we exclude North Dakota and Montana from our analysis.

The Review of Economics and Statistics, October 2018, 100(4): 678–690

doi:10.1162/rest_a_00739

are not driven by two potential confounding factors: the sex composition of the local population (i.e., the adult male/female ratio) and house prices.

For comparison, we revisit the family formation response to the Appalachian coal boom and bust of the 1970s and 1980s. We implement an instrumental variables (IV) strategy similar to that used by Black et al. (2013) that relates birth and marriage outcomes to predicted per capita earnings. We use coal deposits and prices to instrument for earnings in the earlier period and simulated fracking production as an instrument for earnings in the later period. The results of this analysis indicate that the increased earnings associated with the coal boom during the 1970s led to an increase in marriage rates, an increase in the marital birthrate, and a decrease in the nonmarital birthrate. The contrast in findings between periods might suggest that as nonmarital births have become increasingly common, individuals are more likely to respond to increased income with increased fertility, whether or not they are married and not necessarily with an increased likelihood of marriage. We conclude by speculating that social norms play an important role in determining the response of family formation outcomes to economic conditions.

II. Background on Fracking

The exogenous economic shock of fracking production underlying our empirical approach arises from the technological advancements over time in the extraction of shale gas and tight oil, combined with predetermined geological differences across place in fracking potential. For thousands of years, shale plays have trapped deposits of natural gas and oil far below the surface of the earth. Hydraulic fracturing (also referred to as fracing, fracking, hydrofracturing, or hydrofracking) is a well-stimulation technique that involves the high-pressure injection of "fracking fluid" (primarily water, containing sand or other thickening agents) into a well bore to create cracks in the deep-rock formations. This process releases natural gas, petroleum, or brine. This technology has been in use since the 1950s, but it was not economically profitable until two subsequent developments (see Gold, 2014). First, the innovation of horizontal drilling in the 1980s made it possible for wells to be drilled at an angle following layers of fuel deposits rather than to vertically pass through the deposits. Second, experimentation with the fracking fluid formulation in the late 1990s and early 2000s led to cheaper, more cost-effective applications that were capable of splitting shale rock and releasing the oil and gas reserves.

Together these two technologies made oil and gas extraction from shale plays both feasible and economical, resulting in a wave of drilling and production and the widely publicized (in both good and bad terms) fracking boom. In the popular press, this boom has been touted as creating tens of thousands of jobs and providing starting salaries at $50,000 for recent high school graduates, with average earnings in oil and gas between $70,000 and $80,000.[4] Feyrer, Mansur, and Sacerdote (2017) document substantial wage gains associated with local fracking production, including sizable spillover effects to other industries, as well as nearby counties. Their estimates imply that every $1 million of new oil and gas extracted produces $80,000 in wage income, $132,000 in royalty payments and business income, and 0.85 jobs within the county in the year production occurs. Total regional economic impacts are three times larger than the county-specific estimates. They also document that the impacts of new production on wages are persistent, with two-thirds of the wage income increase persisting two years after the initial shock.[5]

Given the employment opportunities created by local fracking booms in both the oil and extraction industries, as well as more broadly in other types of jobs, one would expect that individuals would migrate to fracking areas to take advantage of the potential economic gains. A contemporaneous paper by Wilson (2017) documents a substantial migration response in North Dakota, but much lower rates of migration to fracking counties in the West, South, Midwest, and Northeast. Specifically, he finds that the population of fracking counties in North Dakota increased by 12% to 25%, but by less than 1% in fracking counties in other states. His research suggests that this uneven migration response might reflect uneven information flow.

Researchers have also examined impacts on nonlabor market outcomes. Muehlenbachs, Spiller, and Timmins (2015) estimate hedonic models of property value impacts of shale gas development in Pennsylvania and New York. They find negative impacts on the property valuation of groundwater-dependent homes close to wells, but small positive impacts for piped-water-dependent homes, which they interpret as consistent with benefits from lease payments. Using data from a fracking county in Pennsylvania, Gopalakrishnan and Klaiber (2014) also document heterogeneous impacts of shale gas exploration activity on property values, with a modest reduction in property values for houses within 1 mile of a shale well. A recent working paper by Bartik et al. (2017) estimates that households' willingness to pay for allowing fracking ranges from $1,300 to $1,900 per household annually, reflecting a positive net valuation of improved economic conditions over amenity losses (e.g., increased crime and noise.) Our paper fits into this new, innovative line of research examining the economic and social consequences of the fracking boom.

[4] See, for example, https://www.minneapolisfed.org/publications/fedgazette/desperately-seeking-workers-in-the-oil-patch, http://www.newyorker.com/magazine/2011/04/25/kuwait-on-the-prairie, or http://www.nytimes.com/2012/12/26/us/26montana.html?rref=collection%2Ftimestopic%2FGas%20(Fuel).

[5] Working papers by Allcot and Keniston (2014), Eliason and Timmins (2014), Fetzer (2014), and Maniloff and Mastromanaco (2014) also show large wage gains associated with fracking, on the order of 5% to 24%. A working paper by Cascio and Narayan (2015) suggests that by increasing the wages of less educated men, fracking has led to an increased propensity among high school–age males to drop out of school.

III. Empirical Approach

We exploit cross-sectional, time-series variation in Public Use Microdata Area (PUMA) fracking production to estimate a causal relationship between local economic shocks and subsequent family formation outcomes. Prior to the technological innovations of the early 2000s, the oil and gas deposits extracted from shale plays through fracking were previously unattainable and had no existing economic value. The sudden shock to local economies when fracking came to their area led to increased labor demand, putting upward pressure on wages, including the wages of less educated individuals. We thus use the fracking boom of the late 2000s as an exogenous positive shock to the potential wages of less educated men living in regions that cover a shale play.

Following Feyrer et al. (2017), we capture the extent of the fracking boom in county c by combining geographical information on shale play location with actual oil and gas production to predict variation in production that is solely due to geology and the progression of time. We obtained data on well location and production from a private company, DrillingInfo, through a special use agreement, as we describe in section IV. Because actual production might be correlated with unobservable characteristics related to economic and demographic conditions, we simulate production using only geographic variation in county exposure to a shale play interacted with year effects. The year interaction serves to adjust production amounts for time-varying changes in relevant prices and technology. We estimate equation (1) for all counties over shale plays and then take the exponential of the predicted value:

$$\begin{aligned}
&\ln\left(new\,production_{cy}+1\right)\\
&\quad=\alpha_c+\sum_{\tau=1998}^{2012}\sum_{j=1}^{J}\theta_{\tau j}I\{county\;c\;over\;shale\;play\;j\}\\
&\qquad\times I\{y=\tau\}+\nu_{cy},\\
&sim.\,new\,production_{cy}\\
&\quad=\exp\left(\begin{array}{c}\hat{\alpha}_c+\displaystyle\sum_{\tau=1998}^{2012}\sum_{j=1}^{J}\hat{\theta}_{\tau j}I\{county\;c\;in\;shale\;play\;j\}\\ \times I\{y=\tau\}\end{array}\right)-1.
\end{aligned} \tag{1}$$

The variable $new\ production_{cy}$ is the dollar value of oil and gas production from wells drilled in the current year y located in county c. The main explanatory variables in equation (1) are a set of interactions between an indicator that equals 1 if county c intersects shale play j (measured using ArcGIS software, as described below) and an indicator that equals 1 in year y. This estimates the average impact of being over shale play j on new production and allows this relationship to vary over time as technology and prices change. We then aggregate up from the county/year level to the PUMA/year level.[6] The observed correlation between actual new production and simulated new production is $p = .69$. We divide simulated new production by the PUMA baseline population in 2000 and scale the measure to represent simulated new production per capita in thousands of dollars.[7]

We estimate first-stage effects of simulated new production on the wages of men and women, separately by gender and education. Our baseline specification is not a two-stage least squares model (2SLS) because of concerns about the exclusion restriction (as we describe below), so this is not formally a first-stage model in a 2SLS estimation. The model is described by the following equation:

$$\begin{aligned}
wages_{py}=\alpha_0+\alpha_1\left(simulated\,new\,production_{py}\right)\\
+X'_{py}\xi+\mu_p+\phi_{sy}+\eta_{py}.
\end{aligned} \tag{2}$$

The subscript p refers to PUMA, and y refers to calendar year. The matrix X_{py} is a vector of time-varying PUMA-level controls, which includes the ratio of 18- to 34-year-old women to men, the log average house price, gender-specific shares of 18- to 34-year-old non-Hispanic black, Hispanic, and non-Hispanic other, and gender-specific shares of 18- to 34-year-olds with less than high school, some college, or a four-year college degree. Age-specific sex ratios and race by gender shares are calculated from the Surveillance, Epidemiology, and End Result Program (SEER) population data, which are derived from the U.S. Census population estimates. PUMA-level age-specific gender-by-education shares are aggregated up from county-level estimates in the 2000 Census and the ACS. The housing price data is obtained from the Federal Housing Finance Agency three digit zip code housing price index, which we then link to counties and PUMA and convert to dollars using the median house value from the 2000 Census.

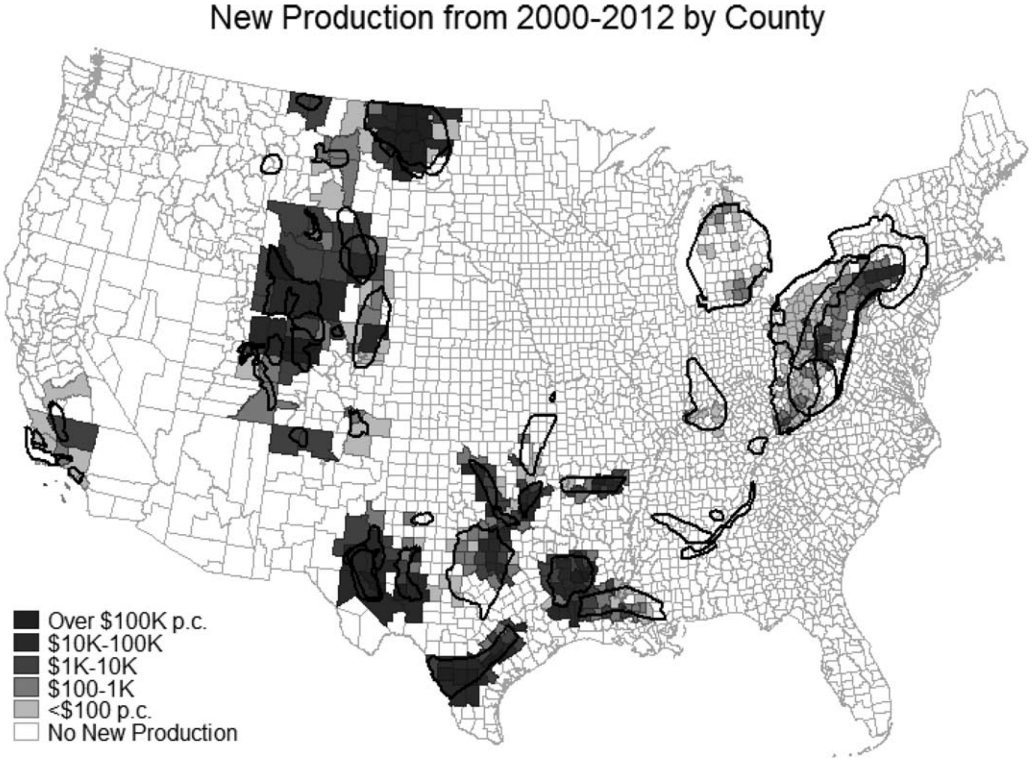
The model also includes controls for time-invariant PUMA effects μ_p and state-specific year effects, ϕ_{sy} to account for fixed differences across PUMA and time trends or shocks in wage outcomes experienced at the state level.[8] The data show a strong, positive relationship between simulated fracking production and the wages of men, and especially non-college-educated men, in the PUMA. We are not

[6] Results are similar if we do not use the exponential to get levels but instead aggregate up the predicted log new production.

[7] We conduct our analysis at the PUMA level rather than the county level primarily because public use Census data on annual level marriage outcomes are not available at areas smaller than the PUMA. Given that previous work has shown that that the labor market impacts of fracking propagate beyond county borders, this is not necessarily disadvantageous.

[8] We do not include PUMA-specific trends in the model because these will (over)control for the response to the shock, an econometric point discussed by Wolfers (2006). When we do estimate the models with PUMA-specific trends included, there is still a statistically significant increase in marital birthrates, albeit the point estimate is much smaller and there is no longer a discernible increase in the nonmarital birthrate. There is still no discernible change in marriage-related outcomes. In table A2 we test alternate specifications that do not include state-by-year effects, which are included in the baseline model to account for geographic spillovers, and find that the results are similar.

FIGURE 1.—TOTAL OIL AND GAS PRODUCTION FROM WELLS DURING FIRST YEAR OF PRODUCTION, 2000–2012



Shale play boundaries are outlined in black.
Source: Constructed by the authors from DrillingInfo.

literally estimating this as a first-stage relationship because our main analysis is a reduced-form analysis. This is an important motivating equation, however, showing that the fracking boom had a first-order effect on male wages.

Figure 1 presents a map showing where shale plays are located and binned total production value from new wells per capita between 2000 and 2012, by county. Production values were uniformly high in North Dakota. Of the sixteen counties in North Dakota with any new production, seven produced $100,000 per capita from new wells between 2000 and 2012, with an average of $244,000 per capita over the entire period. This massive amount of production is due to the large oil reserves in the Bakken shale play. The other locations with the highest levels of simulated production—over $25,000 per capita between 2000 and 2012—include counties in Arkansas, Colorado, Louisiana, Montana, New Mexico, Oklahoma, Pennsylvania, Texas, Utah, and Wyoming. Because of the unique situation in the Bakken region—in particular, the high rates of in-migration—we exclude North Dakota and Montana from our analysis.[9] With that exclusion, there are 88 PUMA with total simulated new production above $25,000 per capita, and an additional 516 with positive total simulated new production below $25,000 per capita.

We estimate the reduced-form relationship between simulated fracking production from new wells and subsequent birth and marriage outcomes using the following specification:

$$\begin{aligned} Y_{py} = {} & \beta_0 + \beta_1 \left(simulated\ new\ production_{py} \right) \\ & + X'_{py}\omega + \mu_p + \phi_{sy} + \varepsilon_{py}. \end{aligned} \tag{3}$$

The key birth and marriage outcome variables of interest (Y_{py}) are defined at the level of a PUMA p and year y, where year denotes year of conception. *Simulated new production*$_{py}$ is measured in thousands of dollars per capita. The matrix X_{py} is the same vector of time-varying PUMA-level controls included in equation (2). Again, the model also includes controls for time-invariant PUMA effects μ_p and state-specific year effects ϕ_{sy}, to account for fixed differences across PUMA and time trends or shocks in birth and marriage outcomes experienced at the state level.

We are ultimately interested in identifying the effect of potential male earnings—an indicator of the marriageability of men, in the conceptual framework of Wilson (1987)—on family formation outcomes.[10] For the reduced-form approach to be informative about the relationship between male economic prospects and family formation outcomes, it

[9] Specifications that include North Dakota and Montana yield a similar pattern of results. There is no effect on marriage outcomes but significant and similarly sized effects on birthrates.

[10] We subsequently use simulated fracking production as an instrument for average earnings, and relate predicted earnings to birth and marriage outcomes.

must be true that simulated fracking production only, or at least primarily, affects subsequent trends in birth and marriage outcomes through its effect on male economic prospects, as captured by measured wages. This condition might not be satisfied if the fracking induced by the existence of shale play affected other factors that might affect family formation, such as house prices.[11] For this reason, our model includes house prices directly as a control variable. If the estimated reduced-form relationship changes with the inclusion of this variable, it would indicate that such variables are potentially driving part of the observed relationship between the local fracking boom and subsequent family formation outcomes. As it turns out, the data do not indicate that to be the case.[12]

IV. Data

This analysis requires detailed information on local-level fracking production, wages, and birth and marriage outcomes. In this section, we provide an overview of the various data sources we draw on for our analysis. More detailed information is available in the data appendix.

A. *Data on Fracking Production*

To construct PUMA-level measures of fracking production, we overlay shale play boundary shapefiles from the Energy Information Administration (EIA) onto U.S. Census Bureau county boundary shapefiles. ArcGIS software is used to assess if counties and shale play intersect.[13] These geographic measures are then combined with well-level quarterly oil and gas production data obtained through a restricted access agreement with DrillingInfo, a private firm that collects lease, permit, and production data on all wells drilled in the United States.[14] The DrillingInfo data file indicates drill date, quarterly production amount, reservoir name, drilling direction (vertical or nonvertical), latitude and longitude, and county. Oil and gas production is reported in barrels and thousands of cubic feet, respectively. We use average annual national prices for oil and gas, recorded by the Energy Information Administration (EIA), to convert production amounts into dollar amounts. All dollar amounts are adjusted to 2010 dollars using the Personal Consumption Expenditure (PCE) price index calculated by the U.S. Bureau of Economic Analysis. We aggregate production values up to the PUMA level using county-to-PUMA mappings from the U.S. Census. In large urban counties that contain multiple PUMA, production values are assigned according to the population share.

B. *Data on Wages*

We use the Quarterly Workforce Indicators (QWI) as our primary source of wage data. The QWI is an aggregation of Longitudinal Employer-Household Dynamics (LEHD) microlevel data collected from unemployment insurance earnings data from participating states and several other sources (U.S. Census, 2014). The QWI is aggregated to the county level and can be tabulated for two-way groups—for example, by gender and age or by gender and education. We construct a PUMA-level group average wage by summing total wage earnings for a given group across all counties in the PUMA and dividing by the total number of jobs in that PUMA for that group.

C. *Data on Birth Outcomes*

We use restricted access Vital Statistics data obtained from the National Center of Health Statistics to construct PUMA-level measures of birth outcomes. These files provide the universe of births between 1997 and 2013 with county identifiers, which we then aggregate to the PUMA level. We date births back to the time of conception by subtracting the length of gestation from the fifteenth of the month of birth. Our analysis sample thus consists of live births that were conceived to women, ages 18 to 34, between 1997 and 2012. We use SEER population estimates to construct age-specific birthrates.[15]

D. *Data on Marriage Outcomes*

We use the 2000 Decennial Census and 2005–2011 American Community Survey (ACS) microdata to construct the PUMA-level share of 18- to 34-year-old women who are never married, married, divorced, and cohabitating (Ruggles et al., 2015). We count women 18 to 34 as cohabitating if they are either the head of the house and an

[11] Kearney and Dettling (2014) and Lovenheim and Mumford (2013) show that an increase in house prices leads to an increase in birthrates among home owners and a decrease among nonowners.

[12] It is possible that rental prices might be more relevant than house prices for this population. County- or PUMA-level rent prices are not available over the entire sample period; however, we are able to construct PUMA-level rent prices from the 2000 Census and 2005–2011 American Community Survey. The estimated effects on this limited sample are slightly smaller but virtually unchanged when we include a control for rental prices.

[13] Special thanks to University of Maryland Geography students Lisa Boland and Michael Bender for their research assistance using ArcGIS software.

[14] These proprietary data are obtained through an academic use agreement with DrillingInfo, available through its academic outreach initiative.

[15] Given our interest in the labor market opportunities and family formation decisions of low-income, less educated men and women, we would ideally examine births to noncollege women specifically. Our baseline analysis focuses on total births to 18- to 34-year-old women largely for data reasons. The nonreporting of maternal education in some state/year cells is a well-known issue with using the natality files for group-level analyses. The NCHS began requiring states to report maternal education according to a 2003 classification starting in 2009. For the twenty states that did not comply with this requirement, maternal education is not recorded in 2009. Starting in 2010, some of these states' education measures were included again. To limit our sample to states with consistent reporting of maternal education severely limits the sample. We thus do not limit the sample by education level in our main sets of analysis, but in a subsequent analysis, we examine birth and marriage outcomes for women separately by education group for the restricted sample, and find the effects are concentrated among noncollege women.

unmarried partner is present or if they are listed as an unmarried partner. The 2005–2011 ACS also includes an indicator that equals 1 if the individual was married in the previous year, which allows us to measure PUMA-level marriage rates. PUMA boundaries changed in 2012, so we restrict our marriage analysis to 2011 and earlier.

E. *Analysis Sample Construction*

Our analysis is estimated at the level of PUMA by year. There are 2,057 total PUMA in the lower 48 states (as compared to 3,109 counties). Two sample restrictions reduce our sample to 2,044 PUMA. First, as noted above, given the unique context of fracking in the Bakken region, in particular the migration response documented in Wilson (2017), we exclude the twelve PUMA in North Dakota and Montana. Second, we exclude Webb County, Texas, because simulated production in this PUMA was over 125% larger than the second-highest-producing PUMA. It is excluded to limit outlier influence. The final sample consists of a balanced panel of 2,044 PUMA observed over the sixteen years from 1997 to 2012.[16]

F. *Summary Statistics*

Table 1 presents PUMA-level summary statistics from the year 2000, before the fracking boom. A "fracking PUMA" is defined as a PUMA with positive simulated production at any point between 2000 and 2012. The statistics reported in this table indicate that fracking areas in our sample are not substantively different from other areas along the dimensions on which this paper is focused. The nonmarital birth share in 2000 for women ages 18 to 34 was 34.2% in nonfracking PUMA and 33.4% in fracking PUMA. Marriage outcomes are also similar across nonfracking and fracking PUMA at baseline, although women ages 18 to 34 in fracking PUMA are slightly more likely to have been married at some point (1.5% less likely to be never married, 1% more likely to be married, 0.3% more likely to be divorced). The share of women ages 18 to 34 married was 44.5% in nonfracking PUMA and 45.5% in fracking PUMA. Overall, labor markets and population characteristics were similar between nonfracking and fracking PUMA, though fracking PUMA had a lower share of college-educated individuals (18.0 versus 15.5 among men and 21.5 versus 18.4 among women, with both differences being statistically significant).

Table 1.—Summary Statistics for Analysis Sample Public Use Microdata Areas, Year 2000 (Prefracking Boom)

	Nonfracking PUMA	Fracking PUMA
Marriage and birth outcomes		
Total births per 1000, women age 18–34	103.9	104.1
Marital births per 1000, women age 18–34	68.6	69.3
Nonmarital births per 1000, women 18–34	35.3	34.7
Percent women 18–34 never married	46.9	45.4
Percent women 18–34 married	44.5	45.5
Percent women 18–34 divorced	5.8	6.1
Percent women 18–34 cohabitating	8.4	8.2
Labor market characteristics		
Average male noncollege earnings (QWI)	43,295	42,930
Average female noncollege earnings (QWI)	27,241	26,560
Average male college earnings (QWI)	79,736	79,803
Average female college earnings (QWI)	45,997	45,436
Population characteristics		
Percent 18–34-year-olds, white-non-Hispanic	62.7	64.1
Percent men 18–34 with college degree	18.0	15.5
Percent women 18–34 with college degree	21.5	18.4
Number of Public Use Microdata Areas (PUMA)	1,440	604

Data on births are from the Vital Statistics natality files. Population counts used in the denominator come from the 2000 Decennial Census. Marital shares come from the 2000 Census. Wage data are obtained from Quarterly Workforce Indicators (QWI). County population characteristics are obtained from U.S. Census Bureau using data from the 2000 Census and SEER population estimates, which are derived from the U.S. Census Bureau's population estimates. See the data appendix for details.

The number of PUMA with active wells increased dramatically over the sample period. The count of PUMA in our analysis sample with active fracking wells in a given year rises from 331 in 2004 to 581 in 2012. Summary statistics about fracking production are reported in appendix table A1. Fracking production is highly skewed across fracking PUMA. For instance, in the year 2012 (the end of our sample period), annual simulated new production for the median fracking PUMA is $3 per capita, but the PUMA at the ninetieth percentile of the production distribution has simulated production of $71 per capita. Annual simulated new production among the top 10% of producing PUMA averaged between $500 and $600 per capita.

IV. Results

A. *Effect of Localized Fracking Boom on Labor Market Outcomes*

We begin our empirical analysis by verifying that local fracking activity led to an increase in potential earnings for non-college-educated men. Table 2 reports the results of the estimation of OLS equation (2), specifying log average earnings in a PUMA/year cell (from the QWI files) as a function of PUMA/year simulated fracking production. There are 29,471 PUMA/year cell observations in this regression, coming from 45 states.[17] The coefficient of

[16] PUMA in our analysis sample with any simulated new production are located in the following 26 states: Alabama, Arkansas, California, Colorado, Georgia, Illinois, Indiana, Kansas, Kentucky, Maryland, Michigan, Mississippi, Missouri, Nebraska, New Jersey, New Mexico, New York, Ohio, Oklahoma, Pennsylvania, Tennessee, Texas, Utah, Virginia, West Virginia, and Wyoming. Eight of these states (Alabama, Georgia, Illinois, Indiana, Maryland, Missouri, New Jersey, and Tennessee) do not have any actual fracking production, but they wind up with predicted fracking production because they have land that overlaps with a shale play. For the PUMA with positive simulated fracking that are in states with no actual production, average simulated production is less than $1 per capita. Recall that our simulation method is intended to create a measure that is a function of geological attributes, not choice variables like local ordinances.

[17] Alaska, Hawaii, Montana, North Dakota, South Dakota, and Massachusetts are excluded. During this period, earnings data from South Dakota and Massachusetts were not available through the QWI.

TABLE 2.—EFFECT OF PUMA-LEVEL SIMULATED NEW PRODUCTION ON LABOR MARKET MEASURES, 1997–2012

	Men			Women		
	Ln(Average Earnings)$_{py}$		Natural Log Jobs/Pop$_{py}$	Ln(Average Earnings)$_{py}$		Natural Log Jobs/Pop$_{py}$
	All Industries (1)	Exclude Oil and Gas Extraction (2)	(3)	All Industries (4)	Exclude Oil and Gas Extraction (5)	(6)
			A. All Workers			
Sim. New Production$_{py}$	0.038***	0.026**	0.052***	0.019***	0.016***	0.012
(thousands of 2010 $ per capita)	(0.013)	(0.011)	(0.020)	(0.006)	(0.005)	(0.010)
Dependent variable mean (in levels)	54,372	54,129	52.53	34,183	34,170	46.86
Observations	29,471	29,471	29,471	29,471	29,471	29,471
			B. Noncollege Workers			
Sim. New Production$_{py}$	0.044***	0.029**	0.048**	0.023***	0.018***	−0.004
(thousands of 2010 $ per capita)	(0.014)	(0.012)	(0.019)	(0.007)	(0.006)	(0.011)
Dependent variable mean (in levels)	43,626	43,365	51.15	28,531	28,518	44.09
Observations	29,471	29,471	29,471	29,471	29,471	29,471
			C. College Workers			
Sim. New Production$_{py}$	0.006	0.001	0.048**	−0.014**	−0.013**	0.038**
(thousands of 2010 $ per capita)	(0.012)	(0.011)	(0.023)	(0.006)	(0.006)	(0.019)
Dependent variable mean (in levels)	80,785	80,586	61.25	48,004	47,991	60.47
Observations	29,466	29,471	29,466	29,469	29,471	29,469

Earnings and job data from the U.S. Census Bureau Quarterly Workforce Indicators file (QWI). Regressions include controls for the female/male 18–34 sex ratio, the natural log of the average house price, gender by race shares for 18–34-year-olds, gender by education shares for 18–34-year-olds, and state × year and PUMA fixed effects and are weighted by the total number of births to 18–34-year-olds in 2000. Standard errors are adjusted for clustering at the PUMA level. ***$p < .01$, **$p < .05$, and *$p < .1$.

interest in column 1 indicates that an additional $1,000 of simulated new production per capita is associated with a statistically significant 3.8% increase in average earnings for men. If we separately look at earnings by educational attainment we see that this increase is concentrated among noncollege men, where an additional $1,000 of simulated new production per capita is associated with a statistically significant 4.4% increase in average earnings.

The increase in earnings is not limited to oil and gas extraction jobs. Column 2 reports that $1,000 per capita of fracking production is associated with a 2.6% increase in average earnings for men in jobs outside those industries, with impacts slightly larger for noncollege men (2.9%). This is consistent with positive spillover effects on other earning opportunities, as documented in Feyrer et al. (2017). In column 3, we see that fracking production is also associated with an increase in the jobs-to-population ratio. One thousand dollars per capita of fracking production is associated with a 5.2% increase in the number of jobs (standard error 0.2). This increase in jobs was experienced by both noncollege and college-educated men. Columns 4, 5, and 6 report analogous results for regressions estimated for women. In earnings specifications, the data indicate similar signed effects, with roughly half the estimated magnitude as for men. However, noncollege women do not observe an increase in jobs, and college-educated women actually observe a small decrease in average earnings. This will be important to keep in mind when we return to interpreting the results.

B. *Effect of Localized Fracking Boom on Birth and Marriage Outcomes*

Having established that local fracking production has a positive effect on the economic prospects of non-college-educated men, we turn to an estimation of the reduced-form relationship between simulated fracking production and birth and marriage outcomes. We start by looking at the impact on the nonmarital birth share, defined as the share of births born to an unmarried mother. A reduction in the nonmarital birth share would be consistent with a reverse marriageable male hypothesis, suggesting that a localized economic shock that raises male earnings is associated with a decrease in the nonmarital birth share. However, to interpret this reduction as a response along those lines, it is necessary to consider what happened to total births and marriage rates. If the number of total births remained constant, a reduction in the nonmarital birth share would reflect a shift from nonmarried births to married births, driven by an increase on the extensive margin of marriage. As shown in table 3, this is not what the data reveal.

Table 3, column 1, indicates that an additional $1,000 of simulated fracking production per capita leads to a statistically insignificant 0.11 percentage point decrease in the nonmarital birth share. The result in column 2 indicates that a localized fracking boom leads to an increase in total births. The point estimate implies that $1,000 of fracking production per capita is associated with an increase of 5.96 births per 1,000 women (standard error of 0.96). In the peak years of the boom, simulated production per capita in the most intensive fracking counties was between $500 and $600 per capita, which would suggest that total births increased by 3 to 3.6 births per 1,000 women, or around 3%. This is consistent with a positive income effect of income on fertility.

Results reported in table 3, columns 3 and 4, show that both marital and nonmarital births increased. The point estimates imply a greater proportional increase in marital births, but the estimated effects are not statistically

TABLE 3.—REDUCED-FORM EFFECT OF SIMULATED NEW PRODUCTION ON PUMA-LEVEL BIRTH OUTCOMES FOR WOMEN 18 TO 34

	Percent Births Nonmarital$_{py}$ (1)	Total Birth Rate$_{py}$ (2)	Marital Birth Rate$_{py}$ (3)	Nonmarital Birth Rate$_{py}$ (4)
Sim. New Production$_{py}$	−0.11	5.96***	3.57***	2.39***
(thousands of 2010 $ per capita)	(0.41)	(0.96)	(1.00)	(0.60)
Dependent means	38.42	101	62.33	38.68
Observations	32,704	32,704	32,704	32,704

Birth data from U.S. Vital Statistics Natality Files. The unit of analysis is the Public Use Microdata Area (PUMA)/year, including 2,044 PUMA from 1997 to 2012. In 2000, approximately 76% of births to women ages 18 to 34 were to women with less than a college degree. Sixty-seven percent of marital births and 94% of nonmarital births were to women with less than a college degree. The impact of simulated new production on marital and nonmarital birthrates is not statistically different. All monetary values are inflation-adjusted to 2010 dollars, using the Personal Consumption Expenditure (PCE) index. All regression models include controls for the natural log average house price, the male/female sex ratio ages 18–34, gender by race shares for 18–34-year-olds, gender by education shares for 18–34-year-olds, and state × year and PUMA fixed effects. Estimates are weighted by the total number of births to 18–34-year-olds in 2000. Standard errors are adjusted for clustering at the PUMA level. ***$p < .01$, **$p < 0.5$, and *$p < .1$.

TABLE 4.—REDUCED-FORM EFFECT OF SIMULATED NEW PRODUCTION ON PUMA-LEVEL MARRIAGE OUTCOMES FOR WOMEN AGES 18 TO 34

	Percent Never Married$_{py}$ (1)	Percent Married$_{py}$ (2)	Percent Newly Married$_{py}$ (3)	Percent Divorced$_{py}$ (4)	Percent Cohabitating$_{py}$ (5)
Sim. New Production$_{py}$	0.06	0.01	0.30	−0.43	0.18
(thousands of 2010 $ per capita)	(1.13)	(0.91)	(0.36)	(0.52)	(0.45)
Dependent mean	56.53	36.03	1.07	4.82	9.19
Observations	16,334	16,334	14,287	16,334	16,334

All outcomes are at the PUMA level and are constructed from the 2000 Decennial Census and the 2005–2011 ACS public use microdata because county is available only for large counties in the public use Census data. The unit of analysis is the PUMA/year, including 2,044 PUMA from 2000 and 2005 to 2011. The share newly married cannot be constructed from the 2000 Decennial Census. All monetary values are inflation adjusted to 2010 dollars, using the PCE index. All regression models include controls for the natural log average house price, the male/female sex ratio ages 18–34, gender by race shares for 18–34-year-olds, gender by education shares for 18–34-year-olds, PUMA and state × year fixed effects. Estimates are weighted by the total number of births to 18–34-year-olds in 2000. Standard errors are adjusted for clustering at the PUMA level. ***$p < .01$, **$p < .05$, and *$p < .1$.

different. The estimated coefficients are, respectively, 3.57 (standard error of 1.0) and 2.39 (standard error of 0.6). Though this is a reduced-form result and not a direct measure of the elasticity of fertility with respect to income, we know of no previous work directly comparing marital and nonmarital birth responses to the same economic shock.

In appendix table A2, we explore the robustness of the estimated birth effects. The response of both marital and nonmarital birthrates is robust to excluding house prices and the sex ratio in the regression model (column 2). It is also robust to estimating the model unweighted (column 3), including year fixed effects rather than state × year (column 4), including shale play by year fixed effects rather than state by year (column 5), and defining the outcome as the natural log of birthrates (column 6).

We also investigate heterogeneous effects by demographic groups. The results are reported in appendix table A3. Column 1 shows that for women ages 35 to 44, there is an increase in the marital birthrate and a decrease in the nonmarital birthrate.[18] The point estimates in columns 2 through 5 show that a sizable increase in both the marital and nonmarital birthrate for non-Hispanic whites, whereas for other race/ethnic groups, the effects are not statistically different from 0. But for those groups, the birth effects are imprecisely measured and racial/ethnic differences cannot be confirmed. The final two columns in the table show that for both the marital and nonmarital birthrates, there is an increase in both first and higher-parity births.

The pattern of results is consistent with a positive income effect on fertility for both married and unmarried couples, but not obviously with a reverse marriageable men hypothesis. To gain more insight into this, we look directly at marriage outcomes. Table 4 reports the results from estimating equation (3) with the dependent variable defined as the percent of women 18 to 34 who are never married, married, divorced, cohabitating, or newly married (married in the previous year). The data give no indication that the economic activity associated with fracking production led to a reduction in the percent never married (column 1) or an increase in the percent married (column 2) or newly married (column 3). Nor do the data give any indication that divorce fell (column 4).[19] Given the rising secular trend in cohabitation, we also test to see if the share of women cohabitating rises with simulated production. We find no evidence of such an effect (column 5).

[18] Estimating this regression for teens shows that the nonmarital birthrate among teens 15 to 17 years old increases by 2.09 (SE 0.84) for an additional $1,000 of production per capita.

[19] The standard bargaining model of marriage in the economics literature posits that as female wages rise relative to male wages, there will be a reduction in marriage because the return to marriage is lower (Becker, 1974). Furthermore, there will be an increase in divorce because the female outside option has increased (Browning et al., 1994). Shenhav (2016) provides empirical support for this prediction for the time periods 1980 and 2010. Exploiting demand shifts in industry/occupation employment cells as an exogenous shock to sex-specific wages, she finds that increases in the relative wage of women led to a decline in the likelihood of marriage for those on the margin of a first marriage. In the fracking context, the male/female wage increased, but the absolute female wage also increased. That leads to offsetting predictions for marriage and divorce outcomes, as the increase in male relative wage should decrease nonmarriage while the increase in female absolute wage should increase nonmarriage. There is also a literature documenting the procyclicality of divorce (Hellerstein & Morrill, 2011; Amato & Beattie, 2011; and Schaller, 2013). In light of that literature, we might have expected a positive effect on divorce rates.

We next consider whether the effect of the localized economic shock on fertility and marriage outcomes varies by social context. In particular, we investigate whether the estimated effects vary with the baseline nonmarital birth share. If a higher rate of nonmarital births is associated with a more accepting social norm than in places where nonmarried births are less common, we might expect to see a larger increase in nonmarital births in response to the positive economic shock associated with fracking in places with higher baseline nonmarital birth shares. As reported in appendix table A4, the data are consistent with this prediction.

In appendix table A4, we report the results of estimating equation (3) separately for counties with a high or low nonmarital birth share, where we define those categories as relative to the median nonmarital birth share among women ages 18 to 34 in the year 2000, which is 33.6% of births. The results indicate a pattern consistent with the social norms prediction. The impacts on marital and nonmarital birthrates are more similar in places where nonmarital births were a larger share of all births and the point estimate on nonmarital birthrates is larger in "high" nonmarital birth share PUMA than in "low" nonmarital birth share PUMA. However, some of these relationships are imprecisely estimated, and we cannot rule out that births responded similarly across "high" and "low" places. To push on this further, we also ran the analysis at the county level (not reported in the table). In that case, the data indicate that the response of nonmarital births in "low" nonmarital birth share counties is close to 0 and significantly different from the marital birth response in those counties.[20] Though not conclusive of social norm effects, these patterns of responses are consistent with the notion that social context partially determines the family formation response to a positive income or earnings shock. We return to this notion when we explicitly compare the findings from the current fracking context to the experience of the Appalachian coal boom and bust of the 1970s and 1980s.

C. *Alternative Specifications*

One possible explanation for the lack of a finding of a marriage effect is that women need time to update their expectations about male economic status or perhaps to observe whether economic improvements are persistent. (Recall that Feyrer et al., 2017, document that the wage increases associated with fracking production show persistence.) To test for this possibility, we specify a regression model where the outcome variable is defined as the difference in the birth or marriage outcome of interest from 2000 and 2011 and estimate that as a function of total simulated production in the PUMA between 2000 and 2011. This definition of fracking production is meant to proxy for the total size of the economic shock during this extended period. This specification is estimated as follows:

$$\begin{aligned}(Y_{p2011} - Y_{p2000}) = \gamma_0 \\ + \gamma_1 \left(\sum_{y=2000}^{2011} \textit{simulated new production}_{py} \right) + X_p^{'}\omega \\ + \phi_s + \varepsilon_p. \qquad (4)\end{aligned}$$

The results of this long-term specification are reported in appendix table A5. As with the annual specification of the regressions, the data indicate a sizable and statistically significant increase in both the marital and nonmarital birthrate. But now the marital birth effect is statistically larger than the nonmarital birth effect. An additional \$1,000 of simulated production per capita increased the marital birth rate by 0.8 births per 1,000 women age 18 to 34 and the nonmarital birthrate by 0.2 births per 1,000 women. There is no discernible change in the percent never married, married, divorced, or cohabiting. Given the weight of the evidence presented thus far, we conclude that there is no evidence to suggest that the economic activity associated with fracking led to an increase in marriage.[21]

Another possible explanation for the lack of evidence of an increase in marriage is that the effect is concentrated among less educated women, and since our analysis sample includes all women ages 18 to 34, we are unable to detect the group-specific changes. Appendix table A6 reports the results of estimating our baseline birth and marriage regressions separately for women ages 18 to 34 with and without a college degree. To do this requires limiting the sample to the subsample of states that record maternal education in the vital statistics natality files for all years. This results in a much smaller sample of 31 states plus the District of Columbia, yielding observations from 1,393 PUMAs. Panel A reports the results for all women estimated on this restricted sample. As found with the full sample in the earlier tables, the results in panel A show statistically significant increases in both the marital and nonmarital birthrates. Panel B reports results for non-college-educated women. There are sizable increases in the marital and nonmarital birthrates, and still no effect on marriage outcomes. For college-educated women, there is only a small, positive effect on the marital birthrate, a small negative effect on the

[20] In specifications not reported in the table, we also estimate the model with an interaction term between simulated new production and a continuous variable measuring year 2000 county-level nonmarital birth share. That specification yielded a statistically insignificant coefficient on the interaction term of interest. In addition, we estimated the model with an interaction of simulated new production and year 2000 measure of the percent of the population who are religious adherents, as measured from the Association of Religion Data Archives county-level church membership. This variable captures the total percent of the population recorded as members by Christian, Jewish, Islamic, or Eastern religious institutions but does not necessarily capture the level of engagement. That interaction term did not enter the model with statistical significance for any of the birth or marriage outcomes.

[21] Alternative specifications using lagged new production or dynamic AR(1) processes to identify long-run impacts also provide no evidence of long-run increases on marriage outcomes.

TABLE 5.—COMPARISON TO THE APPALACHIAN COAL BOOM CONTEXT: IV ESTIMATES OF THE EFFECT OF EARNINGS ON MARITAL BIRTHS, NONMARITAL BIRTHS, AND SHARE OF WOMEN NEVER MARRIED

	A. Appalachian Coal Boom (1969–1987)			
	Women 18–34			Women 15–34
Natural Log of	Percent Births Nonmarital (1)	Marital Birthrate (2)	Nonmarital Birthrate (3)	Percent Never Married (4)
Ln earnings per Capita$_{py}$	−3.03***	0.75***	−2.55***	−0.96***
	(0.55)	(0.23)	(0.47)	(0.18)
First-stage F-statistic	56.3	56.3	56.3	34.0
Dependent means (in levels)	13.94	81.63	12.81	51.69
Observations	2,546	2,546	2,546	268
	B. Fracking Oil and Gas Boom (1997–2012)			
	Women 18–34			Women 15–34
Natural Log of	Percent Births Nonmarital	Marital Birthrate	Nonmarital Birthrate	Percent Never Married
Ln earnings Per Capita$_{py}$	−0.11	1.24***	1.24***	0.17
	(0.25)	(0.43)	(0.43)	(0.21)
First-stage F-statistic	11.8	11.8	11.8	8.6
Dependent means (in levels)	38.42	62.33	38.68	59.82
Observations	32,704	32,704	32,704	4,088

All birth data from Vital Statistics. Share never married is calculated from the 1970 and 1980 censuses (provided by National Historical Geographic Information System), as well as the 2000 Decennial Census and 2011 ACS. Earnings per capita from both periods are constructed from the BEA. We use coal reserve values to instrument for the natural log of earnings per capita for the coal boom, similar to Black et al. (2013). Data on coal reserves for 1969–1988 are constructed using data provided by Black et al. (2013). The sample for the coal boom period includes 134 Public Use Microdata Areas (PUMA). The sample for the fracking boom period includes 2,044 PUMA. All regression models include controls for the female population 18–34, male/female sex ratio ages 18–34, gender by race shares, and state × year and county fixed effects. In panel B we also control for the natural log average house price. This measure is not available for the earlier period. Estimates are weighted by the total number of births to 18–34-year-olds in 1970 and 2000, respectively. Standard errors are adjusted for clustering at the PUMA level. ***$p < .01$, **$p < .05$, and *$p < .1$.

nonmarital birthrate, and no statistically significant effects on marriage outcomes.

D. *Comparison to the Effects of the Appalachian Coal Boom in the 1970s and 1980s*

In a paper titled "Are Children Normal?" Black et al. (2013) analyze how fertility among married couples responded to the increase in male earnings in the Appalachian coal-mining region of the United States in the mid-1970s. The Appalachian coal boom began in the 1970s when energy prices spiked and continued through the 1980s until energy prices plummeted. Black et al. (2013) exploit variation in county-level coal reserves and yearly energy prices to estimate the causal effect of changes in earnings on marital birthrates. Specifically, they estimate an IV regression of the marital birthrate in a county as a function of total county earnings, where county earnings are predicted as a function of the value of coal reserves in the county. (They estimate the model in first differences.) They motivate and interpret their analysis as an empirical test of Becker's (1960) contention that children are normal goods, meaning that the demand for children increases in response to rising income. Black et al. estimate that a 10% increase in county earnings associated with the coal boom leads to a 7% increase in the marital birthrate.

During the time period Black et al., studied nonmarital birthrates were much less common than they are today. It is thus interesting to consider whether the nonmarital birth response to the coal boom might have been different from what we are observing in the context of the localized fracking boom. In table 5, we revisit the fertility response to the Appalachian coal boom by looking at both married and unmarried births. This extension builds directly on the reduced-form results reported in appendix table A4 showing that places with low nonmarital birth shares experienced a larger increase in marital births than nonmarital births.

We follow the approach of Black et al. (2013) in estimating an IV regression of birthrates as a function of predicted log earnings. We use coal reserve and price measures provided by Black et al. to instrument for the natural log of earnings per capita during the coal boom period. During the later period, we estimate an identical specification but use simulated fracking production as an instrument for log earnings. To be sure that earnings measures are comparable across periods, we use the natural log of PUMA-level per capita earnings from the U.S. Bureau of Economic Analysis (BEA) for both periods, since QWI wage data are not available before 1990.[22]

The estimates reported in table 5, panel A, column 2, imply that a 10% increase in earnings led to a 7.5% increase in the marital birthrate among women ages 18 to 34. This is

[22] The IV specification requires the exclusion restriction assumption that the only channel through which a localized fracking boom affected marriage and birth outcomes is through earnings. Two obvious threats to this assumption are changes in house prices and changes in sex ratios. But recall that these variables are controlled for in our model, so the IV estimate is conditional on those variables. Furthermore, results are unchanged when those variables are excluded. Other potential threats to the validity of the exclusion restriction are changes in unearned income, changes in female earnings as opposed to male earnings, and changes in the male-to-female wage ratio. Conceptually, all of these variables could have a direct effect on marriage and birth outcomes and are affected by local fracking booms. Without additional instruments, we are unable to adequately control for these confounding factors. The estimated effects from this IV analysis should thus be interpreted with caution.

comparable in magnitude to the estimated relationship during the fracking boom, as reported in panel B, column 2. The IV estimate of the relationship between per capita earnings associated with fracking and marital births implies that a 10% increase in earnings led to a 12.4% increase in the marital birthrate.

The nonmarital birth response is very different between contexts. As reported in table 5, panel A, column 3, the estimated relationship between earnings and nonmarried births during the Appalachian coal boom is negative and significant. A 10% increase in earnings associated with the coal boom led to a 25.5% reduction in the nonmarital birthrate. Off of a mean of 12.8, this represents a reduction of 3.3 nonmarital births per 1,000 women. This contrasts sharply with the estimated relationship identified from variation in fracking production in the 2000s. As reported in panel B, the IV results imply that a 10% increase in earnings associated with fracking production led to a 12.4% increase in nonmarital births.

The data also indicate a very different marriage response in the earlier and later periods. Column 4 reports the results of estimating the IV model for the dependent variable "share of women age 15–34 never married." The data suggest that in the earlier period, a 10% increase in per capita earnings was associated with a 9.6% reduction in the share of women age 15 to 34 who were never married. We use this modified age range because in the 1970 census, there are two age groups available: 15–24 and 25–34. The point estimate from the later period implies a statistically insignificant 1.7% increase for a 10% increase in per capita earnings.[23]

The contrast of findings between the context of the cool boom and bust of the 1970s and 1980s and the fracking boom of the 2000s is consistent with the notion that social context matters. In the earlier period, when nonmarital births were still far from the norm, couples responded to the increase in earnings with increased rates of marriage and increased marital births but no increase in births outside marriage. In the later period, both marital and nonmarital births increased significantly in response to the positive economic shock. And unlike during the Appalachian coal boom, there is no discernible increase in marriage in response to the positive local economic shock associated with fracking.

Although this evidence is consistent with changing social norms, it is not definitive. There are other potential explanations for the differential responses observed between periods. For instance, the earnings impact of the coal boom and bust was particular to male earnings (per Black et al., 2013). But as we saw in table 2, fracking increased the potential earnings of women as well, albeit to a lesser extent than for men. An increase in female earnings could mute the positive effect of male earnings on marriage rates, leading to the null effect found in the later period, and also make it more financially feasible for an unmarried woman to have a child without a spouse. It might also be the case that fracking jobs are particularly onerous, so that an increase in earnings for men directly employed in fracking-related jobs would make them less desirable marriage partners than would an increase in earnings for men employed in other jobs. Another possibility is that the migration response was significantly different between the two periods such that more of the men in fracking counties (even with the exclusion of North Dakota and Montana) are likely to be transient and hence less likely to be ideal marriage partners.[24] The coal boom also incorporates both a boom and a bust, which might yield different results if improving and declining male employment prospects have asymmetric impacts. However, if we restrict our analysis of the Appalachian region to the boom period (1969–1982), the resulting point estimates show a similar pattern. Specifically, the marital birthrates increase by 0.36% (SE 0.20) and nonmarital birthrates decrease by 2.48% (SE 0.52), leading to a lower nonmarital birth share. Still, we cannot rule out the role of factors other than social norms, and we thus view the comparisons and contrasts of results between panels A and B in table 5 as interesting and our interpretation of them necessarily speculative.

V. Conclusion

The fracking boom of the post-2005 period led to sizable improvements in the earnings potential of non-college-educated men in regions located over geological shale plays. We use this context as a rare opportunity to investigate whether an increase in the employment and earnings potential of less educated men leads to an increase in marriage rates and a corresponding decrease in nonmarital births. Our analysis is motivated by an interest in testing whether the reverse of the marriageable men hypothesis holds: As the economic prospects of less educated men improve, are couples more likely to marry and are women less likely to have a child outside a marital union? The data suggest that at least in the short term, if we extrapolate from the experience of the fracking boom, the answer is likely no.

[23] To make a more direct comparison, we estimate these regressions for a common set of Appalachian PUMA in both periods. Unfortunately sample size limitations preclude a useful analysis along these lines. For the three Appalachian states from the Black et al. (2013) sample, the point estimates on predicted log earnings due to fracking production are as follows: 0.54 (SE 0.37) for the marital birthrate, 1.25 (SE 0.63) for the nonmarital birthrate, and 0.03 (SE 0.59) for the female share never married. These results are imprecisely estimated but consistent with an increase in nonmarital births and no increase in marriage during this more recent context.

[24] Though we cannot rule out this possibility, a comparison of the migration response between episodes does not indicate qualitatively different migration responses. Black, McKinnish, and Sanders (2005) document a sizable male migration response to the coal boom. They report that the county population of men ages 2 to 29 during the boom years of 1970 to 1980 increased by 0.09 log points, with a standard error of .040. By way of comparison, we find that the population of men ages 20 to 29 increased by 0.063 log points among the top 20% most productive fracking counties in our sample between 2000 and 2011.

The results of our analysis suggest that local-area fracking production led to an increase in both marital and nonmarital births and no increase in marriage rates. The finding of a positive birth response is consistent with a positive income effect on births as found in the previous literature. We build on this previous finding by separately examining the response of marital and nonmarital births to the same shock. The data suggest that the marital birth response is larger than the nonmarital birth response in areas with relatively low nonmarital birth shares at baseline, while the marital and nonmarital responses are similar in areas with high nonmarital birth shares at baseline. This would be consistent with a role of social norms in driving the family formation response to a local economic shock.

To further investigate the possibility that social context matters, we have compared the family formation response to the fracking boom of the 2000s to the family formation response to the Appalachian coal boom and bust of the 1970s and 1980s. These are similar economic shocks in generally similar types of places. The data indicate that the increased earnings associated with the coal boom during those earlier decades led to an increase in marriage rates and marital births and no increase in nonmarital births. In contrast, the increase in earnings associated with fracking in more recent years led to an increase in both marital and nonmarital births and no increase in marriage rates. This contrast is consistent with the notion that the family formation response to economic circumstances depends on social context.

In conclusion, we find no evidence from the fracking context to support the proposition that as the economic prospects of less educated men improve, couples are more likely to marry before having children. An important caveat to broadly interpreting this finding is that this experience might not be generalizable to other types of economic improvements. For instance, if there were to be a boom in manufacturing or technology sector jobs for non-college-educated men today, the resulting family formation effect might be different. The results of our analysis, as with most reduced-form empirical analyses, are context specific, and one must therefore be cautious in using them to draw general conclusions.

Furthermore, our findings do not imply that the decline in the economic position of men in certain communities and demographic groups over the past four decades has not been a primary driver of the increase in nonmarital birthrates and decrease in marriage rates in these communities and groups. It is quite possible that the reduction in male marriageability among less educated men in earlier decades was the driving force that led to a decline in marriage and a corresponding rise in nonmarital childbearing, but now that nonmarital childbearing has become so commonplace, a new social norm has been set and an increase in male economic prospects does not have the same effect as it would have in a different time or place. In other words, economics might have led to a new social structure such that we are now in a new paradigm. The proposition that individuals respond to economic circumstances in ways that are shaped by prevailing social norms warrants further empirical examination.

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