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WHO NEEDS A FRACKING EDUCATION? THE EDUCATIONAL RESPONSE TO LOW-SKILL BIASED TECHNOLOGICAL CHANGE

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ABSTRACT

We explore the educational response to fracking, a recent technological breakthrough in the oil and gas industry, taking advantage of the timing of its diffusion and spatial variation in shale reserves. We show that fracking has significantly increased relative demand for less-educated male labor and high school dropout rates of male teens, both overall and relative to females. Our estimates imply that, absent fracking, the teen male dropout rate would have been 1 percentage point lower over 2011-15 in the average labor market with shale reserves, implying an elasticity of school enrollment with respect to earnings below historical estimates. Fracking increased earnings more among young men than teenage boys, suggesting that educational decisions respond to improved earnings prospects, not just opportunity costs. Other explanations for our findings, like changes in school quality, migration, or demographics, receive less empirical support.

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I. Introduction

Technological change, much of it complementary to skilled labor, has defined the U.S. economy for more than a century. Its effects have not been confined to labor markets. By increasing the return to education, skill-biased technological change (SBTC) spurred stunning growth in educational attainment through cohorts born in the mid-20th century (Goldin and Katz, 2008). In more recent generations, however, increases in the relative demand for skill have consistently outstripped increases in its relative supply, suggesting that the elasticity of educational attainment with respect to the skill premium may now be low and contributing to rising wage inequality (Katz and Murphy, 1992; Goldin and Katz, 2008). Yet, producing credible micro-level evidence on how SBTC affects educational investment decisions is difficult given the typically widespread nature of technological change.

This paper estimates the educational response to a recent technological breakthrough in a specific industry – oil and gas extraction. By pumping large quantities of fluids at high pressure down a wellbore into horizontal wells in a target rock formation, hydraulic fracturing – or "fracking" – has made it possible to extract oil and natural gas from shale plays unreachable through conventional technologies (U.S. Environmental Protection Agency, 2013). Recent research suggests that the local employment impacts of fracking have been sizable and extend beyond oil and gas extraction, expanding industries like mining, transportation, and construction that disproportionately employ less-educated men (Feyrer, Mansur, and Sacerdote, 2017).[1] Additionally, labor markets that have tightened due to fracking have seen employers in other sectors cut education and experience requirements to fill positions (Modestino, Shoag, and Balance, 2017). There is also direct evidence that labor demand shocks from fracking have

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[1] For general employment effects, see also Bartik et al. (forthcoming) and Maniloff and Mastromonaco (2017). Krupnick and Echarte (2017) provide a recent review of the broader literature.

favored men without a college degree (Bartik, 2018; Kearney and Wilson, 2018). We go further, documenting that fracking has improved the labor market prospects of male high school dropouts by more than any other group. Fracking thus provides a case study in "less" skill-biased technological change, with implications for education that are the reverse of the standard SBTC story: if responsive to the skill premium, educational investments should have fallen due to fracking.

We explore the evolution of educational outcomes across areas with different shale oil and gas endowments as fracking has spread.[2] Because of the identification challenges posed by the migratory response to fracking, we focus on high school enrollment and dropout decisions of teenagers, measured in both survey and administrative data (the Census/American Community Survey (ACS) and the Common Core of Data (CCD), respectively). Our empirical approach is to compare local labor markets – commuting zones (CZs) – with different shale oil and gas reserves over time.[3] We focus on 14 states with major shale plays,[4] and our preferred models are demanding, removing bias from time-varying shocks to enrollment that vary across states and across CZs with different pre-fracking observables. Like Bartik et al. (forthcoming), our models also allow fracking's impacts to phase-in gradually. We assume that they begin to unfold across

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[2] Our approach is methodologically similar to that of Michaels (2011), which estimates the long-term consequences of oil abundance in the U.S. South, including impacts on educational attainment in the resident adult population.

[3] CZs are collections of contiguous counties (possibly crossing state boundaries) that were strongly linked on the basis of commuting patterns in the 1990s (Tolbert and Sizer, 1996). Like metropolitan areas, CZs have been used in past research to define local labor markets (e.g., Autor and Dorn, 2013; Autor, Dorn, and Hanson, 2013a, 2013b; Chetty et al., 2014), but they have the relative advantage of covering the entire United States, including rural areas. CZs are thus ideal for our analysis, since fracking is largely a rural phenomenon. Feyrer, Mansur, and Sacerdote (2017) also find that CZs do a good job of capturing the local economic impacts of fracking, which do not respect county boundaries.

[4] Shale plays are shale formations with similar geologic and geographic properties that have significant quantities of natural gas. We define major plays as those shale plays that have reserve estimates reported by Energy Information Administration. See Data Appendix.

the country in 2006, but our substantive conclusions are robust to incorporating geographic variation in timing and to allowing fracking opportunities to arise nationwide at an earlier date.

We find that fracking has slowed the rate of decline in high school dropout among male teenagers since the early 2000s. Our estimates imply that, due to fracking, the dropout rate of 17-18-year-old boys (as measured in the Census/ACS) was 1.1 percentage points higher and the ratio of 11^{th} and 12^{th} grade enrollment (as measured in the CCD) to the 17-18-year-old population 1.4 percentage points lower over 2011-15 in the average community with shale gas and/or oil reserves. This substantive conclusion is robust to how we estimate reserves and, as noted, to changes in timing the onset of fracking's diffusion. We also show in the Census/ACS that the estimates are accounted for largely by boys who have not recently migrated, suggesting that fracking has changed educational decisions among teenagers, rather than their residential choices. Controlling for changes in the demographics of teen boys and allowing for increases in compulsory schooling ages to have disparate effects in CZs with larger shale endowments also influence these findings very little. Estimates for girls are indistinguishable from zero, but we can often rule out effects as large as we find for boys, consistent with the incidence of fracking's labor market impacts.

These estimates are reduced-form, however, leaving causal mechanisms uncertain. Labor demand shocks from fracking could have encouraged boys to drop out not just by increasing the longer-term earnings prospects of male high school dropouts, but also by raising the short-term opportunity cost of staying in school. While we cannot rule out a role for increased opportunity costs, we show that fracking improved labor market outcomes significantly more for young men than teenage boys, suggesting a change in the perceived return to schooling contributed to the male dropout response. Even so, perceptions of this return based on older cohorts may differ from the actual return for affected cohorts, since fracking may affect that return through other channels – by reducing school quality, for instance. Though recent research has found male schooling choices to be more sensitive to school quality (Autor et al., forthcoming), we can rule out even small effects of fracking on overall per-pupil school spending and average class sizes, suggesting that male schooling choices would have to be considerably more sensitive to quality measures orthogonal to school resources to explain our findings.

These conclusions diverge from those of Weber (2014), who finds that natural gas fracking in four states has increased the share of the adult population with a high school degree and reduced the dropout share. The difference may owe to migration: fracking has generated modest, but consistent net in-migration of working-age adults (Bartik, 2018; Wilson, 2017). By focusing on a younger, less mobile population, we attempt to isolate educational decisions from residential ones, like previous studies of resource booms and busts (Black, McKinnish, and Sanders, 2005; Emery, Ferrer, and Green, 2012; Morissette, Chan, and Lu, 2015), and as noted, our survey data allow us to rule out migration as a confounder. By incorporating more states and multiple data sources, our analysis is also more comprehensive than concurrent studies of fracking's impacts on the schooling decisions of young people (e.g., Marchand and Weber, 2015; Rickman, Wang, and Winters, 2017; Zuo, Schieffer, and Buck, 2018). More states allow for consideration of the mediating effects of changes in state education policies like compulsory schooling laws. Multiple data sources also allow us to closely examine which mechanisms are driving the dropout response.[5]

[5] Exploiting school district variation in shale geology across the state of Texas and temporal variation in energy prices over the 2000s, Marchand and Weber (2015) find declines in vocational and economically disadvantaged student enrollment. Rickman, Wang, and Winters (2017) focus on the educational attainment of native-born individuals aged 18-24 using cross-state variation between three states and a synthetic control group. Zuo, Schieffer, and Buck (2018) provide a complementary analysis using different sources of fracking-related variation and focusing on aggregate high school enrollment from the CCD, without separate estimates by sex.

More broadly, we add to a recent surge of papers on the educational impacts of aggregate economic shocks, which has considered settings from trade shocks in Mexico and around the world (Atkin, 2016; Blanchard and Olney, 2017), to the housing price bubble in the U.S. (Charles, Hurst, and Notowidigdo, 2018), to infrastructure and workfare programs in India (Adukia, Asher, and Novosad, 2017; Shah and Steinberg, 2017). Like these studies, we exploit localized variation in the incidence of an economic shock. To our knowledge, however, this paper is among the first to present micro-level evidence of the downstream effects of technological change on schooling decisions.

A back-of-the-envelope calculation based on our estimates yields an elasticity of male high school enrollment with respect to adult male earnings of around 0.18, below the lower bound of the range of elasticities estimated by Black, McKinnish, and Sanders (2005) in their study of the 1970s coal boom and 1980s coal bust. We might have expected if anything a stronger response to technological change in resource extraction than to even long-lived resource price shocks. Yet, these findings are consistent with a low supply elasticity of educated labor today (Goldin and Katz, 2008), and provide some of the first credible micro-level evidence that a weak educational response to SBTC may be contributing to widening wage inequality.

II. Background on Fracking

A. Geography and Timing

Figure 1 plots geographic variation in fracking potential for the 17 states containing at least one CZ that is part of a major shale play.[6] We derive the reserve measure from the 2011 map of shale plays published by the Energy Information Administration (EIA) and the maximum

[6] We focus on all shale plays with oil reserves and/or gas reserves reported by the EIA for at least four years between 2008 and 2015. There are 17 states with at least one CZ containing a major shale play by this definition: Arkansas, Colorado, Kentucky, Louisiana, Maryland, Montana, Nebraska, New Mexico, New York, North Dakota, Ohio, Oklahoma, Pennsylvania, Texas, Virginia, West Virginia, and Wyoming.

EIA-reported economically recoverable oil and gas reserves by major shale play across 2008 to 2015.[7] Overlaying the shale map to CZs, we allocate oil and gas reserves to CZs based on the fraction of each play that they represent, then convert them to a common metric that captures the amount of heating energy that they contain – millions of British Thermal Units (MMBTUs).

There is considerable regional variation in reserves per capita (2000 population), with clusters of high-reserve areas in the Western, Southern, and mid-Atlantic regions. These areas lie atop different major shale plays: the Bakken (in Montana and North Dakota), the Barnett, Eagle Ford, Fayetteville, Haynesville-Bossier, and Woodford (in Louisiana, Oklahoma, and Texas), and the Marcellus (in Pennsylvania and West Virginia). However, although the reserve measure is rather blunt, constructed from just the intersection of a CZ with a major play and the overall play's economic potential,[8] there is also variation in the magnitude of per-capita reserves across areas that are close geographically. This means we can identify the effects of fracking from within-state variation in reserves across CZs, an approach that helpfully sweeps out the shared effects of other state-level shocks, such as changes in state education policy or other aggregate economic developments. It also mitigates the influence of outliers in the reserve distribution, which are concentrated in states like North Dakota and Texas.

Our estimation sample restricts attention to the 14 states in Figure 1 where data on employment and earnings by sex are available from the year 2000 forward from the Quarterly

[7] Economically recoverable reserves are estimated volumes of hydrocarbon resources that analysis of geologic and engineering data demonstrates with reasonable certainty are recoverable under existing economic and operating conditions.

[8] We therefore don't take advantage of within-play differences in "prospectivity," which could reflect local efforts to identify economically recoverable reserves and be independently related to trends in educational attainment. See the Data Appendix for a complete description of how the reserve measure was calculated. We explore the robustness of our estimates to different ways of constructing reserves in Table 5.

Workforce Indicators (QWI), our primary source on the labor market impacts of fracking.[9] We also trim the sample to exclude the smallest 5% and largest 10% of CZs within each state, based on population in the year 2000.[10] Table 1 Panel A shows that, while the average CZ in this sample has substantial shale reserves, its shale oil and gas *production* as of 2000 was not that high. This is expected: shale gas and oil reserves have become exploitable only as horizontal drilling and fracking have spread. For our main analysis, we date the start of widespread use of fracking to 2006 – a year that predates the first frack dates of the highest-reserve plays but marks the rough beginning of the application of the technology in lower-reserve plays.[11] The aggregate annual production trends shown in Figure 2 for horizontal or directional ("unconventional") wells, based on data from DrillingInfo, are consistent with shale gas and oil production taking off after 2005.[12]

B. Evidence on the Spread of Fracking

The combination of geography and time forms the core of our identification strategy: if fracking has increased the propensity of teens to drop out of school, dropout rates should have

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[9] We lose Arkansas, Kentucky, and Wyoming due to missing QWI data for the year 2000. Schooling estimates are substantively similar when we include these states, as shown in the Appendix. Data sources are discussed below and described completely in the Data Appendix.

[10] The motivation for this is twofold: our outcome measures are especially noisy for the smallest CZs, and the largest CZs often include large cities that do not seem a valid counterfactual for the typical CZ with shale reserves. We drop a total of 57 CZs due to this sample restriction. Our findings for schooling attainment are substantively similar if we drop the smallest 10%, rather than 5%, of CZs within each state, but are more sensitive to including larger cities. See Appendix.

[11] We prefer this approach to one taking advantage of variation in first frack dates across shale plays given the potential endogeneity of play-specific timing and the fact a lack of ACS data for 2001-2004 makes variation in timing more difficult to exploit convincingly. However, our estimates are qualitatively similar when we exploit this timing (Table 5). Bartik et al. (forthcoming) report first frack dates of 2008 for both the Marcellus play and the Haynesville-Bossier play, for shale gas, and of 2007 and 2009, respectively, for the Bakken play and the Eagle Ford play, for shale oil. Smaller plays, like Avalon Bone-Spring (oil), Fayetteville (gas), Woodford (gas), were reportedly first fracked in 2005 or 2006.

[12] The data are annual aggregates of monthly well-level production data from DrillingInfo. Following prior research using these data, we classify production from horizontal and directional ("unconventional") wells as fracking, or as coming from shale. To be conservative, we classify unknown well types as vertical (or "conventional") wells. In Table 5, we assess the robustness of our findings to the choice of 2006 in various ways.

increased more – or declined less, given that dropout rates were declining over this period (Murnane, 2013) – as fracking has spread, and more so in places with larger shale endowments. It is therefore useful to begin our investigation by establishing that our reserves measure and assumptions about timing predict changes in local economic activity consistent with existing estimates. Because it is important for the interpretation of the dropout findings that follow in Section III, we also document the sex and skill bias in these changes to local economic activity using data from the QWI, which aggregates administrative microdata on jobs and earnings covering 95% of private sector workers (see Data Appendix) and has recently been used in other work to expose granularity in the local labor market impacts of fracking (Kearney and Wilson, 2018).

Figure 3 Panel A presents event-study estimates of the impacts of fracking on oil and gas production per capita (in thousands of MMBTUs), based on CZ-by-year aggregates of the DrillingInfo data underlying Figure 2. More specifically, the figure plots estimates of the θ_τ's from the following model:

(1) $$y_{zst} = \sum_{\tau \neq 2005} \theta_\tau reserves_z D_t^\tau + \lambda_{st} + \delta_z + \varepsilon_{zst},$$

where y_{zst} represents per-capita oil and gas production in CZ z in state s in year t; $reserves_z$ is the CZ's predicted per-capita shale reserves (in MMBTUs; Figure 1); D_t^τ represents a year dummy set to one when t=τ, and λ_{st} and δ_z are vectors of state-by-year and CZ fixed effects, respectively. The θ_τ's trace out what happened over time to the within-state slope that characterizes the relationship between production and $reserves_z$; as fracking has spread, for instance, the $reserves_z$ gradient for shale production should have gone from flat (Table 1 Panel A) to upward-sloping. The capped vertical lines represent 90% confidence intervals on these estimates, with standard errors clustered on CZ.

As expected, there is no evidence of an impact of shale reserves on conventional oil and gas production from vertical wells: the slope on reserves does not significantly change between 2005 (the year for the omitted interaction) and any subsequent (or prior) years. However, the coefficient estimates for production from *horizontal* wells (fracking) suggest that production picked up in higher-reserve CZs several years after 2005; the first statistically significant change in slope, relative to 2005, actually does not occur until 2011. But the overall pattern is not surprising given the trends in Figure 2, and it is reassuring that greater growth occurs in CZs with higher shale oil and gas endowments per our reserve measure.

Wells needed to be drilled before they can produce, however, so the labor demand effects of fracking should have been felt before large production impacts. And it is arguably these employment shocks – rather than shale production *per se* – that are more salient for schooling choices. The black triangles in Figure 3 Panel B represent event-study estimates for the jobs-to-population ratio for men ages 25 and over.[13] The underlying data are CZ-by-year aggregates of quarterly county- and sex-level data from the QWI, for the numerator, and annual sex-specific Census-based estimates of county population from SEER (Survey, Epidemiology, and End Results Program at the National Cancer Institute), for the denominator (see Data Appendix). The event-study coefficients imply that, despite significantly lower jobs-to-population levels in 2000 (Table 1, Panel B), higher-reserve CZs did not experience different job growth for men between 2000 and 2005. Thereafter, however, higher-reserve CZs start gaining jobs for men faster than lower-reserve CZs in the same state. This phenomenon strengthens through the end of the period

[13] We focus on the entire population ages 25 and over because the educational breakdown of jobs and monthly earnings by sex, explored in Table 3 and Figure 4, is available only for this broad age group in the QWI. See the Data Appendix for a complete description of the QWI data. We consider impacts on jobs and earnings in narrower age bands, across all education categories, in the section on causal mechanisms (Table 7).

with a distinct increase in effect size between 2010 and 2012. The female coefficients (gray diamonds) are much lower in magnitude, implying smaller impacts on female employment.

Another way of measuring the labor demand shocks from fracking is to estimate its impacts on the earnings prospects. To this end, we follow Charles, Hurst, and Notowidigdo (2018) in estimating impacts on the natural log of "expected" monthly earnings, with expected monthly earnings defined as the product of the average monthly earnings of adults ages 25 and over times the jobs-to-population ratio. As shown in Figure 3 Panel C, this measure suggests more sizable labor market impacts over 2006 to 2010, but a difference between an immediate and a later post-fracking period is still noticeable.

To characterize the pattern of labor market impacts revealed by model (1) parsimoniously, and to quantify the patterns shown, we consider a slightly modified difference-in-differences model, with two post-fracking periods:

(2) $$y_{zst} = \theta_1 reserves_z D_t^{06-10} + \theta_2 reserves_z D_t^{11-15} + \lambda_{st} + \delta_z + \varepsilon_{zst},$$

where D_t^{06-10} and D_t^{11-15} are dummies set to one when *t* is in the ranges 2006-10 and 2011-15, respectively.[14] θ_1 thus represents the average change in the within-state $reserves_z$ gradient between 2006-10 and 2000-05, and θ_2 represents the change between 2011-15 and 2000-05. To validate this specification, we also test whether estimates of θ_1 and θ_2 are different.

Unless otherwise noted, throughout the remainder of the paper, we estimate model (2) including time-varying effects of the pre-existing (year 2000) CZ observables in Table 1 Panel D. As shown in column 2, CZs with higher per-capita shale reserves differ in some ways from

[14] We have also estimated deviation-from-trend models, which effectively impose separate linear fits on the event-study coefficients before and after 2005. The identifying assumption is that, in the absence of fracking, any trending in the $reserves_z$ gradient in the pre-period would have simply continued. Such a model allows some of the effects of fracking to be immediate (post-2005 $reserves_z$ intercept shift) and some to accumulate over time (post-2005 change in the trend in the $reserves_z$ gradient). The results are quite similar to what we report for model (2).

those with lower reserves in the same state: they had significantly lower median annual household incomes and black population shares in 2000, for example.[15] Though the coefficients are not too large in magnitude,[16] failure to allow for time-varying effects of these characteristics, in addition to time-varying effects of $reserves_z$, could therefore bias our estimates.

The first two columns of Table 2 Panel A present estimates of model (2) for the jobs-to-population ratio of people ages 25 and over (x100), separately by sex, including this vector of additional controls.[17] Column 3 then shows estimates of the difference in the male and female coefficients. Panel B repeats this exercise for the natural log of expected monthly earnings. In each panel, the first row uses a simple linear transformation to convert $\hat{\theta}_2$ into a more interpretable magnitude: the implied effects of fracking over 2011-15 for the average CZ with any reserves ($\hat{\theta}_2 \times \hat{\mu}_{reserves|reserves>0}$, where $\hat{\mu}_{reserves|reserves>0}$ = 42,060 MMBTUs, from Table 1 Panel A). The average CZ with reserves gained more jobs for men than women due to fracking: about 43 male jobs for every 1000 adult men and about 6 female jobs for every 1000 adult women by 2011-15. It also saw substantial earnings growth: expected earnings were 10.4% higher for men, and 2.6% higher for women, by 2011-15. These estimates are roughly in line with other nationwide studies examining the labor market effects of fracking.[18] For both

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[15] Each CZ is given equal weight, except in the case of outcomes constructed from public-use microdata, in which case we weight for efficiency purposes by cell size. For the purposes of the fixed effects, we assign CZs to the states in which the majority of their population resides.

[16] Table 1 column 3 shows that the predicted difference in median household income in the average CZ with shale reserves versus without is about $886, which represents 2.7% of the variable's mean (figure in italics) and 15.4% of its standard deviation. We arrive at similar calculations regarding magnitudes for other variables where there are statistically significant coefficients on $reserves_z$.

[17] Appendix Table A1 provides the corresponding estimates for oil and gas production, measured both in thousands of MMBTUs and in millions of 2012 dollars. Estimates of $\theta_1$and θ_2 for unconventional production are relatively less precise, consistent with Figure 3 Panel A. Effects for the conventional production measures are not statistically significant.

[18] For example, Bartik et al. (forthcoming) estimate an increase in employment of 3.6% to 5.4% and earnings of 4.4% to 6.9% in the top quartile of "fracking potential" counties. Extensive reviews of the relevant literature provided by Maniloff and Mastromonaco (2017) and Krupnick and Echarte (2017) find employment effects that range from 1.3% to 14.4% and earnings effects that range from 1.8% to 16.7%.

outcomes, there is a significant difference in effects between the immediate and later post-fracking periods, and men were significantly more affected.

As evidence on whether there is indeed a *skill* bias to these labor demand shocks, in addition to a bias toward men, Table 3 breaks out the jobs and earnings gains by four categories of educational attainment. We estimate the sex- and education-specific CZ population by multiplying the overall adult CZ population by the share of adult men or women ages 25-64 with the relevant completed education in the Census or ACS.[19] For brevity, we show only implied effects by 2011-15 for the average CZ with any reserves, calculated as earlier described.[20] Figure 4 provides a graphical representation of the corresponding event-study (model (1)) estimates.

Men of all education levels have experienced job and expected earnings gains as a result of fracking (Panel A). However, gains along both dimensions have been largest among high school dropouts: by 2011-15, the average male dropout in a CZ with any reserves had a 7 percent higher chance of holding a job due to fracking – a figure marginally significantly larger than the gain for high school graduates (p=0.063). In addition, male dropouts ages 25 and over in the average CZ with reserves would have expected to earn, across 2011-15, 12.7% more than they otherwise would have. This is not significantly greater than the earnings growth expected by high school graduates (p=0.185), but it is significantly greater than the expected earnings growth among men with some college or more. Among women, jobs gains were more similar across education categories, but earnings gains were also most substantial at the bottom of the distribution (Panel B).

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[19] We estimate this time-, sex-, and education-varying population share using the 2000 Census and the 2005-2015 ACS. We linearly interpolate these shares for 2001 through 2004.

[20] Appendix Tables A2 and A3 show the corresponding coefficient estimates from model (2) for both outcomes along with p-values on the difference in coefficients between 2006-10 and 2011-15. We reject equality at the 0.05 significance level in all cases except for the jobs-to-population ratio for female dropouts (p=0.055).

III. Fracking and High School Dropout

We have established that our measure of shale reserves and assumptions about the timing of the spread of fracking predict changes in local economic activity consistent with existing findings. We have also added to existing understanding of these changes by exploiting more granular data: not only did places with more shale reserves experience more growth in labor demand starting around 2006 – particularly in 2011 and later – those increases in labor demand were weighted toward men, and among men, those less-educated. Male high school dropouts saw the greatest improvements in their jobs and earnings prospects.

Collectively, these findings suggest that the educational decisions of young men, including on the high school dropout margin, are likely to have been affected by fracking. Our next goal is to estimate the reduced-form effect of fracking on high school dropout by sex, using the same models that we employed in Section II; we defer further discussion of causal mechanisms to Section IV.

A. Data

We use two sources of information on dropout; neither is ideal, but they provide complementary evidence on the phenomenon of interest.

First, we construct high school dropout rates from the public-use microdata samples (PUMS) of the 2000 Census and the 2005 through 2015 ACS (Ruggles et al., 2015). These data provide person-level information on sex, age, school enrollment, and educational attainment and identify local geography down to the Public Use Microdata Area (PUMA) level. While PUMAs are not the same as CZs, they can be allocated to CZs based on the division of county population across PUMAs and the mapping between CZs and counties. As with our analysis on the spread of fracking, we choose to aggregate to the CZ level to capture the local economic impacts of fracking, which do not respect county or PUMA boundaries (Feyrer, Mansur, and Sacerdote, 2017). We focus on 17- and 18-year-olds – an age group for which dropout decisions are salient but migration for work is low – and define dropout as having not recently been enrolled in school and not having a high school degree or GED (see Data Appendix).

Though suitable for this analysis, the Census and ACS have limitations. First, small samples (5% in the Census and 1% in the ACS) contribute to imprecision. Second, no information on PUMA is available in the 2001 to 2004 ACS, limiting the years available to establish common outcomes trends in the pre-fracking period, which are implicit in a difference-in-differences specification. Third, the dropout rates from the Census and ACS are not necessarily comparable due to differences in survey timing.[21]

Our second data source – the CCD – helps to address these limitations. The CCD includes annual school-district level data on high school enrollment by grade and sex. Enrollment data in the CCD are in principle consistently measured and cover all school districts in the United States. After accounting for some lapses in coverage, aggregating 11th and 12th grade enrollment (combined) from the district to the CZ level, and normalizing by estimates of the 17-18-year-old population from SEER, we have an annual series on sex-specific enrollment-to-population ratios beginning in the 1999-00 academic year (represented as 2000 to follow). The main limitation of this series is that data are missing for 2009-10 (2010).[22]

[21] The Census, officially collected on April 1, solicits school enrollment as of February 1. By contrast, the ACS is fielded throughout the year (with survey month not publicly reported) and asks about school enrollment over the past three months. Because a new school year begins typically in August or September, a higher share of the 17- and 18-year-olds surveyed in the fall – and therefore a higher share of 17- and 18-year-olds in the ACS – will be of age to be enrolled in 12th grade or beyond. If teens sometimes make dropout decisions based on grade rather than age, and if that tendency happens to be correlated with $reserves_z$, our estimates from the Census/ACS analysis could be biased.

[22] See Data Appendix for more detail on both sources.

B. Baseline Estimates

Figure 5 presents event-study estimates of model (1) for the 11th and 12th grade enrollment-to-population ratio (Panel A) and for the high school dropout rate of 17- and 18-year-olds (Panel B). In both cases, we restrict attention to the same sample of CZs considered in Section II; the standard errors underlying the 90% confidence intervals (capped vertical lines) are clustered on CZ; and we include time-varying effects of the CZ observables in Table 1 Panel D, in an attempt to improve precision. We also weight the estimates in Panel B by cell size.

The coefficients are noisier than those in Figure 3 Panel B and Figure 4. However, the estimates are consistent with fracking weakening the attachment of teenage boys to high school. As was the case in Figures 3 and 4 for jobs-to-population ratios and expected earnings, the gradient between male dropout and $reserves_z$ (black triangles) shows a distinctly different trend after 2005 from the flat trend that preceded it. Starting in 2006, higher-reserve CZs begin to see larger reductions – or smaller increases – in male enrollment-to-population ratios relative to lower-reserve CZs in the same state (Panel A). For dropout, the pattern is flipped, showing relative increases in dropout in high versus low-reserve CZs in 2006 and later (Panel B). In both panels, the point estimates appear consistent with a larger impact, on average, in 2011 and later, precisely when fracking appears to start having larger labor market impacts. The event-study coefficients for female dropout rates (gray diamonds) do not display this pattern.

Table 4 presents estimates from the more restrictive difference-in-differences specification, model (2), for teen male (column 1) and female (column 2) enrollment-to-population ratios (Panel A) and dropout rates (Panel B). Regressions in both panels include state-by-year and CZ fixed effects as well as time-varying effects of year 2000 CZ observables, to match the figures. Considering first the estimates for boys, neither outcome variable experiences a statistically significant change in the $reserves_z$ gradient between 2006-10 and 2000-05, though coefficients move in the direction of increased dropout. However, there is a significantly greater decrease (increase) in the enrollment-to-population ratio (dropout rate) for relatively high reserve CZs between 2011-15 and 2000-05, and the data reject equality of the difference-in-differences coefficients for 2006-10 and 2011-15. The point estimates imply that, due to fracking, male enrollment-to-population (dropout rates) in the average CZ with any reserves were 1.4 percentage points lower (1.12 percentage points higher) over 2011-15. Column 2 confirms that a similar phenomenon is not happening for teen girls, and column 3 shows the differences in estimated effects across sex are statistically significant.

C. Specification Checks

Table 5 explores the sensitivity of our estimates to a number of changes in the econometric specification. For reference, Panel A repeats the sex-specific estimates for both enrollment-to-population and dropout rates from Table 4. Columns 5 and 6 of each panel demonstrate how the changes in specification affect sex-specific estimates of the impact of fracking on the natural log of expected monthly earnings (Table 2). This is a useful benchmark, since specifications that generate weaker "first stage" impacts on the labor market should also generate weaker impacts on school enrollment, if changes in the labor market are the key mechanism linking fracking to schooling decisions. Throughout, we give predicted effects by 2011-15 in the average CZ with any reserves.[23]

To begin, we consider sensitivity to controls. In Panel B, we first drop the controls for time-varying effects of the year 2000 CZ observables in Table 1 Panel D. The impacts on the

[23] Appendix Table A4, A5, and A6 show the corresponding coefficients from the specifications in Panels B, C, and D, respectively. Plots of event-study coefficients for each of these specifications are in Appendix Figures A1, A2, and A3.

estimates are generally small and not consistent across outcomes. However, substituting year fixed effects only for the state-by-year fixed effects in our preferred model slightly raises the magnitude of the estimated impacts for boys. Such a finding may arise if fracking opportunities in a CZ also increased dropout propensities of boys elsewhere in the state. But an alternative explanation is that higher-reserve CZs are in states where dropout rates would have been falling less quickly over time for other reasons. Although the estimates are largely similar, state-by-year fixed effects help to ensure our estimates are not contaminated by these state-specific shocks. The final specification in Panel B provides unweighted estimates for the dropout outcome. The standard errors actually fall slightly in the unweighted specification, suggesting the potential importance of small outliers. The point estimates, however, remain basically unchanged.

Panel C considers different assumptions about timing. First, we assign each CZ with reserves the first frack year for the earliest-fracked shale play it lies atop (t_z^*), with dates as reported by Bartik et al. (forthcoming). The model of interest is then:

(3) $$y_{zst} = \theta\, reserves_z post_{zt} + \lambda_{st} + \delta_z + \varepsilon_{zst},$$

where $post_{zt} = 1[t \geq t_z^*]$. The intuition of the empirical approach still applies; the only change is the introduction of variation across states in when the $reserves_z$ gradient should begin to change. For both schooling outcomes, the estimates are now smaller in magnitude for boys. However, estimated impacts on adult male earnings are as well, and implied effects of local earnings growth on schooling decisions for boys are quite similar to what we saw at baseline, particularly for the enrollment-to-population ratio. Nevertheless, we can no longer reject equality of the impacts on dropout rates across boys and girls. The next specification is a simple difference-in-differences, comparing 2011-15 to the year 2000 only, and so acknowledging that fracking occurred earlier in some places. The primary consequence is to reduce statistical power.

Panel D explores the robustness of our conclusions to alternative predictions of local shale reserves. The first two approaches weight oil and gas reserves by price rather than energy content in the aggregate reserve prediction under extreme assumptions: using the peak ratio of oil to gas prices (from 2012) and the peak ratio of gas to oil prices (from 2003).[24] We then apply the earliest available shale play reserve estimates (from 2008), combining oil and gas into common energy units as in our original measure.[25] For boys, the implied effects of earnings growth on enrollment (column 1) and dropout (column 3) are weakened the most when using 2008 reserve estimates, but they are still evident, suggesting that using more recent reserves data mainly serves to improve statistical power. Dropout estimates for girls continue to remain indistinguishable from zero, and in most cases remain significantly lower in magnitude than those for boys. Overall, the implications are the same regardless of how reserves are measured: fracking appears to have lowered schooling attainment for teenage boys much more so than for girls.[26]

D. Education, Location, or Policy?

Collectively, the specification checks above produce estimates that are substantively similar to those at baseline: the schooling decisions of teenage boys appear to have been

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[24] In essence, we bound the relative weights placed on oil and gas when combined using prices; although there is temporal and geographic variation in energy prices, price expectations as relevant for our analysis would arguably fall between these extremes of annual industry-standard price benchmarks.

[25] Shale oil reserve estimates are not available until 2011, so our local per-capita reserve estimates for 2008 are imputed for oil with 2011 reserve values.

[26] Appendix Table A7 shows sensitivity of our schooling estimates to changes in the estimation sample. The reduced-form estimates are slightly smaller in magnitude when we include CZs from the three states with major shale plays but without QWI data for 2000, though we still reject equality of effects by 2011-15 across sex. Dropping the bottom 10% rather than the bottom 5% of a state's CZs (based on size) has lesser effect on the estimates than dropping the top 5% rather than the top 10% of a state's CZs. In the latter case, the estimated effect on male dropout shrinks (but remains significant) and we can no longer reject equality of effects with females. We think our preferred sample selection rule is justifiable given that large CZs are unlikely to provide a valid counterfactual for smaller CZs where fracking is more prevalent, and indeed, the balance tests in Table 1 column 2 look less compelling with larger CZs included.

significantly more negatively affected by fracking than those of teenage girls. Before we move on to *why* educational choices were affected by fracking, it is useful to confirm that the estimates indeed reflect educational decisions rather than residential ones. Concerns over migration motivated our study of teenagers, but could migration still be affecting our estimates? Setting aside the question of migration, other explanations for our findings besides fracking, like heterogeneity in responses to changes in state education policies, are also possible.

Consider migration first, which we can observe in Census and ACS microdata. Because of the endogeneity of migration, we do not condition estimation on it but rather decompose the outcome variable into a dummy for being *both* a dropout *and* a recent migrant (in the past year in the ACS and in the past 5 years in the Census) and a dummy for being *both* a dropout *and not* a recent migrant; the sum of coefficients across these two outcomes is equal to the total baseline dropout effect. As shown in Table 6 Panel B, only the estimates for being both a dropout *and not* a recent migrant are statistically significant for boys and statistically different between boys and girls. The vast majority of the reduced-form estimate thus appears to be accounted for by educational choices rather than residential ones.

Even if not driven by migration, shifts in the demographics of the local population could be incidentally correlated with the spread of fracking and could therefore influence our estimates. In the first set of estimates in Panel C, we control for time-varying local shares of the Census and ACS 17- and 18-year-olds who are black, Hispanic, or who have recently migrated. The implied effects on dropout change little, suggesting that fracking is not strongly correlated with changes in the composition of 17- and 18-year-olds, at least on these observable dimensions.[27]

[27] We would like to include family background measures, like family income or parental education, but these variables are only available for the selected sample of 17- and 18-year-olds in the Census and ACS who still live with their parents.

As a final check that our estimates reflect a response to fracking rather than to other changes in the local environment, we control for changes in the age at which states allow children to drop out of high school. Common effects of such changes are of course already captured in our state-by-year fixed effects. But suppose that dropout rates in higher-reserve areas fell relatively less in response to the increases in compulsory schooling ages that occurred over our estimation period.[28] Then a dropout effect we are attributing to fracking may really be caused by heterogeneity in the impacts of state education policy. In the last rows of Table 6, Panel C, we assess this possibility directly by including as a control the interaction between $reserves_z$ and a state-by-time-varying indicator for the requirement that a person be enrolled in school until age 17 or 18. We do this both for the dropout rate of 17-18-year-olds and for the enrollment-to-population ratio. For males, the estimates get slightly larger in magnitude, though not significantly so, suggesting that such heterogeneity either is working against us seeing an effect or is difficult to detect given the generally small changes in dropout in response to recent changes in compulsory schooling laws (Oreopoulos, 2009).

IV. Mechanisms: Theory and Evidence

The dropout findings capture the reduced-form effects of fracking – the effects of fracking on schooling decisions working through any channel. While such estimates are independently interesting, we also care about causal mechanisms. For example, in Becker's classic (1964) model of human capital investment, candidate mechanisms for an increase in dropout rates would include more than just a reduction in the return to a high school degree;

[28] Over the period of interest, the only states in our estimation sample where compulsory schooling laws changed to require enrollment until age 17 or age 18 were Colorado, Maryland, Nebraska, and West Virginia. We obtained compulsory schooling ages from 2000 forward from NCES at https://nces.ed.gov/programs/digest/d15/tables/dt15_234.10.asp?current=yes, and assumed any change occurred in the earliest possible year over the relevant interval.

increases in the costs of staying in high school, namely increases in the opportunity cost, matter as well. But fracking can also affect the return to a high school degree by influencing school quality, which may rise due to additional funding, or fall due to crowding.[29]

To formalize the discussion of channels, consider a choice between two states – high school graduation and high school dropout – among 18-year-olds. Those who choose to graduate enter the labor market one year later than those who do not. The decision point is normalized to year 0, so that dropouts enter the labor market in year 0, and graduates enter in year 1; all individuals then work through year T in discrete time indexed by t. Individuals i vary in their ability, d_i, which is uniformly distributed over [0,1]. For simplicity, assume that earnings in both states are known at the time of the decision, with $Y_{HS}(t) > Y_{noHS}(t)\ \forall t$. Also assume that the indirect or psychic costs of remaining in high school until period 1, $\varphi(d_i)$, are decreasing in ability, so that $\varphi' < 0$. Most potential dropouts are enrolled in public school, so direct costs of remaining enrolled are zero (no tuition).

Teen i chooses high school graduation if the lifetime payoff from graduation exceeds that from dropout, i.e., if $V_{HS}(d_i) > V_{noHS}$. With ability uniformly distributed on [0,1], the break-even ability for staying in school, d^*, is also the high school dropout rate. If psychic costs are then linearly decreasing in ability, with $\varphi(d_i) = \varphi \times (1 - d_i)$, $\varphi > 0$, d^* takes on the intuitive closed-form expression:

$$d^* = 1 - \frac{1}{\varphi}\left(\underbrace{\sum_{t=1}^{T} \frac{1}{(1+r)^t}\big(Y_{HS}(t) - Y_{noHS}(t)\big)}_{return\ to\ high\ school} - \underbrace{Y_{noHS}(0)}_{opportunity\ cost}\right).$$

[29] Increased family incomes could be another channel through which fracking influences the return to a degree. For example, children from families whose incomes have risen due to fracking may be gaining more skills from high school than they would have otherwise. To the extent that responses would be similar for girls and boys, they would not help to explain the gap in estimated effects by sex, which is our focal point here.

Thus, the dropout rate is a decreasing function of the return to high school – or earnings prospects of dropouts over their working lives – and an increasing function of the opportunity cost of staying in school.

How does this map to the evidence that we've presented thus far? We've shown that the teen male dropout rate has risen and the return to education among adult men has fallen by more over time in areas with more shale reserves. Evidence of the same phenomena for females is much weaker, suggesting that the reduction in the return is the key causal pathway for our findings. But an increase in the opportunity cost of school enrollment for boys could also be contributing to our estimates. This would especially be the case if boys were present-biased (Cadena and Keys, 2015; Lavecchia, Lu, and Oreopoulos, 2016) and fracking increased earnings of teenage boys. The standard human capital model of course assumes that that they would be forward looking, which might not be realistic.

In an attempt to disentangle the relevance of the opportunity cost and returns to education channels, Table 7 considers the jobs-to-population ratio (Panel A) and natural log of expected monthly earnings (Panel B) for various narrow age groups and separately by sex, calculated from QWI data and population estimates as earlier described.[30] For each outcome, we provide the year 2000 mean (column 1), difference-in-differences estimates from model (2) with additional controls (columns 2 to 3), the p-value on equality of the two coefficient estimates between year groups (column 4), sexes (for 2011-15, column 5), and age groups (for 2011-15, column 6), and the implied effect of fracking by 2011-15 for the average CZ with any reserves (column 7). Though fracking led only to a small increase in the likelihood of 14-18-year-old boys having

[30] Unfortunately, information on jobs and earnings by educational attainment *and* sex is not available over narrow age ranges considered in the table. See Appendix Figure A5 for event-study representations of the effects for these outcomes.

jobs (Panel A), it substantially increased the wages and/or hours of those with jobs (Panel B). Those estimates are nevertheless significantly smaller than those observed for only slightly older males, ages 19-24 and 25-34, whose experiences on the job market are arguably particularly salient for prospective dropouts in forming expectations of the future.

We of course cannot rule out that teenagers are misinformed, over-reacting to a misplaced notion that well-paying jobs are immediately available to them that are not. But the estimates are also consistent with changes in the return to education playing a role in the dropout decision of boys. Nevertheless, changes in the return to education *among adult men* may not represent the changes in the *expected* return across the lifecycle for the cohorts of interest. The fracking boom could have reduced school quality, for example, lowering the skill (and future productivity) gains associated obtaining a high school degree.

Table 8 considers whether school quality has declined as a result of fracking, potentially lowering the return to schooling independently of labor market developments. We would have liked to measure school quality with some measure of school output, like test scores, but available test score data are not geographically disaggregated enough or do not span enough years to apply our empirical approach.[31] Instead, we consider class size and per-pupil school spending and revenues, using CZ aggregates of annual data from the CCD and the Census of Governments and Annual Surveys of State and Local Government Finances, respectively.

[31] The National Assessment of Educational Progress (NAEP) spans the period of interest, but reports data only for states and selected large school districts. With passage of the No Child Left Behind Act, states were required to test children in grades 3 through 8 in math and reading and publicly report the test results starting in 2002-03. Unfortunately, the tests and reporting practices differ across states, so comprehensive estimates are unavailable for our study period.

The first three rows of the table show that there was no effect of fracking on average class size, overall per-pupil spending, or overall per-pupil revenue.[32] Given the standard errors, we can rule out small contributions of changes in school resources to our schooling estimates.[33] Fracking did affect the *composition* of revenues, though, consistent with expectations. In line with fracking increasing property values (Bartik et al. forthcoming), fracking appears to have increased local (property-tax based) revenues per student enrolled in public schools.[34] However, this positive effect on local revenues has been offset by negative revenue effects at the federal and state levels.[35]

While speculative, these auxiliary findings suggest that labor market developments associated with fracking are important to our school enrollment findings, and that teenage boys have based their decisions on more than just opportunity costs. Incorporating all causal pathways, our estimates yield an elasticity of high school enrollment with respect to adult

[32] See Appendix Figure A6 for the corresponding event-study estimates. The revenue and spending results are in line with the null effects obtained in other national-scale analysis (Bartik et al., forthcoming) but by construction are unable to capture the heterogeneous effects uncovered in regional analyses (Marchand and Weber, 2015; Newell and Raimi, 2015; Ratledge and Zachary, 2017; Weber, Burnett, and Xiarchos, 2016, Zuo, Schieffer, and Buck, 2018). Our class size results obtained at the national level also cannot produce the heterogeneity documented in regional analyses (Marchand and Weber, 2015; Ratledge and Zachary, 2017).

[33] For example, the upper bound on the 95% confidence interval for the average class size effect of fracking in a CZ with any reserves is less than a one student increase. As a point of comparison, Dynarski, Hyman, and Schanzenbach (2013) find that being randomly assigned to attend a small class in kindergarten (with 13 to 17 students) instead of a regular-sized class (with 22 to 25 students) – so having 5 to 12 fewer students in the same class – increased the probability of attending college by 2.7 percentage points, with an effect size for males 1.6 percentage points higher than that for females. The implied effect of an additional student on the male-female difference in college attendance is thus about 0.2 percentage point (1.6/8.5, where 8.5 is the average class size increase). If we adopted the same 0.2 percentage point implied effect of an additional student for high school dropout, we would be able to account for less than 20% of the increase in the male-female gap in dropout rates.

[34] Not all studies find positive effects of fracking on property values. Muehlenbachs, Spiller, and Timmins (2015) find large negative impacts on nearby groundwater-dependent house prices, though at a broader geographic scale, they find positive impacts that diminish over time. Gopalakrishnan and Klaiber (2013) also find negative impacts on property values.

[35] State governments often redistribute local tax revenue across school districts in an effort to narrow spending differences between more and less property-wealthy districts. Relatedly, the federal government's primary grants program (Title I) is distributed on the basis of child poverty. If fracking has reduced child poverty rates, any educational impacts for local students may have eventually been offset by reductions in Title I funds.

earnings of 0.18.[36] This elasticity estimate is on the lower end of the range of elasticities that Black, McKinnish, and Sanders (2005) estimate by exploiting earnings variation from the coal boom and bust of the 1970s and 1980s, which suggest that school enrollment falls by 2.2% to 7.2% for every 10% increase in earnings. The difference in truly comparable elasticities between the two studies would, moreover, arguably be larger: the variation in Black, McKinnish, and Sanders (2005) arises from long-lived, but ultimately transitory, price-shocks, biasing downward their elasticity estimate as would be comparable to our setting of permanent, technological change. The effect size for young men in our study is thus small by historical standards.

V. Conclusion

Over the past decade, the widespread diffusion of horizontal drilling and fracking has fueled a structural transformation of local economies across the United States – from Pennsylvania to North Dakota – increasing local incomes and helping to set the U.S. on a path toward energy independence. Using high-frequency outcomes data and taking advantage of geographic and temporal variation in fracking, we have demonstrated that this structural transformation has had the additional consequence of slowing reductions in high school dropout rates among teenagers, particularly the young males whose longer-term labor market prospects it has more greatly affected. Though we cannot completely rule out other causal pathways, such as an increased opportunity cost, we marshal additional evidence to support the conclusion that perceptions of a reduced longer-term return to a high school degree for men were an important causal mechanism. For example, fracking did not appreciably increase jobs for teenagers.

[36] We calculate this elasticity by comparing the 1.43 percentage point decrease in the teen male enrollment-to-population ratio from our baseline specification – which amounts to roughly a 1.9% decline given the baseline ratio (Table 1 Panel C) – to a 10.4% earnings gain from fracking for adult males (Table 3).

Fracking also appears to have changed school resources too little for changes in school quality to be an important factor in our findings.

Do these findings provide cause for concern? As we describe, the decision to drop out of school could well be a rational one in the face of increases in later-life job opportunities for dropouts. Nevertheless, some students could be making mistakes in dropping out. There are also social benefits from completing high school that are ignored in private dropout decisions (e.g., Lochner and Moretti, 2004; Dee, 2004; Milligan, Moretti, and Oreopolous, 2004). Fracking may thus be generating sub-optimally low levels of education among some individuals who would already likely be relatively low-skilled, with possible implications for future productivity and the social safety net. Whether the human capital of a generation of young men has been permanently affected is an open question, one that cannot be readily answered with our data and research design.[37]

Second, and more broadly, we present new evidence on the relationship between technology and educational attainment. With fracking, we have a technology that complements low-skilled labor and one whose use is geographically constrained in a way that allows for credible identification of its impacts on educational attainment. We find evidence of reductions in educational attainment at the bottom of the skill distribution, and that longer-term declines in the return to education could be an important contributing factor. In addition, converting our point estimates to an elasticity and comparing our findings directly with earlier work, we find our estimated effect sizes to be consistent with a low supply elasticity of educated labor today (Goldin and Katz, 2008).

[37] Although those students who drop out may re-enroll later in their lives (Emery, Ferrer, and Green (2012) do find some evidence for such re-enrollment in British Columbia), we cannot investigate this possibility with Census data.

Altogether, our study suggests a smaller response in educational attainment to skill-biased labor demand shocks today than estimated in earlier decades and supports the view that SBTC has contributed to wage inequality. Even so, the effects of technological change on education may be heterogeneous across the skill distribution. Future understanding of this relationship would therefore benefit from exploration of other episodes of localized technological change, particularly ones favoring the highly skilled.

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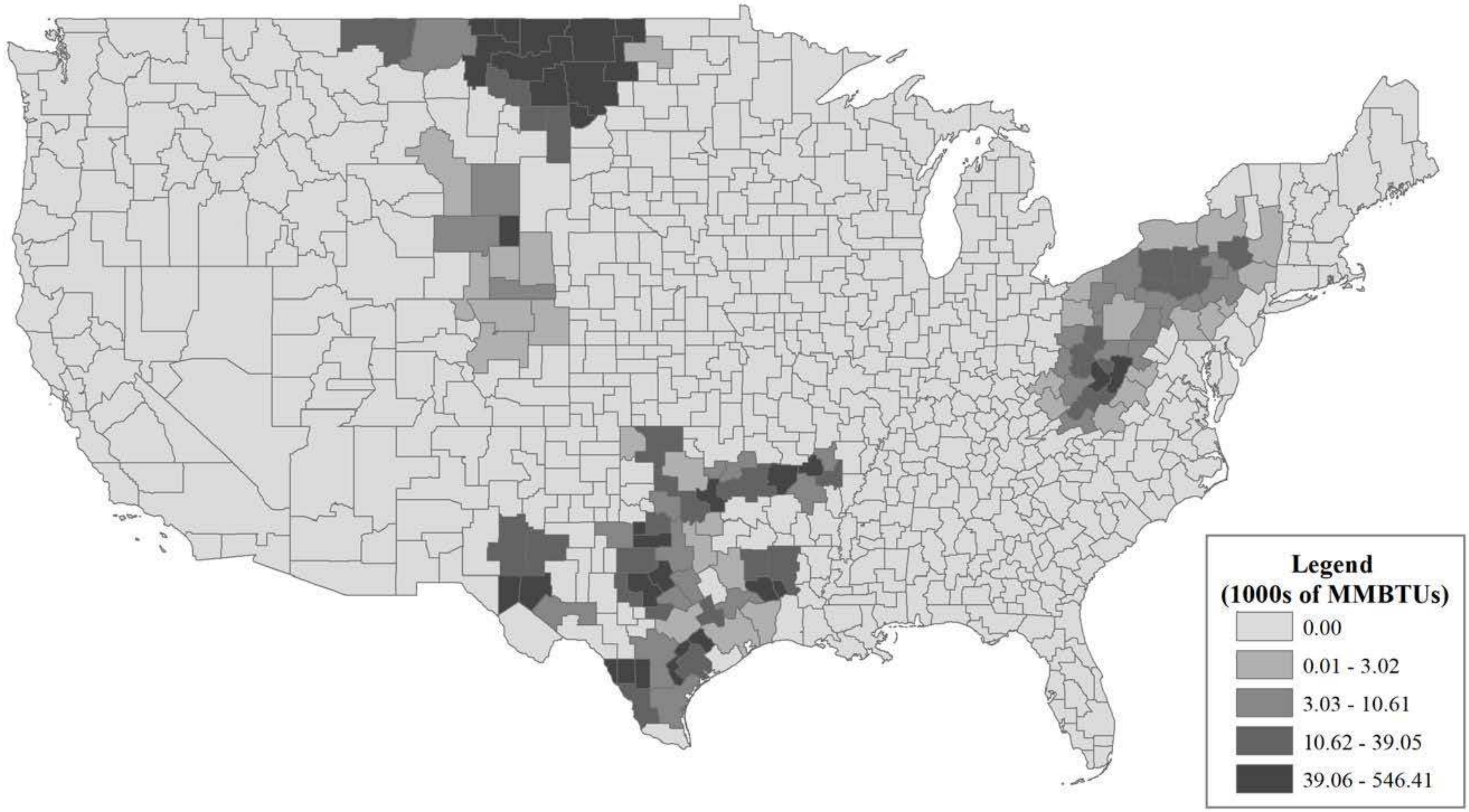
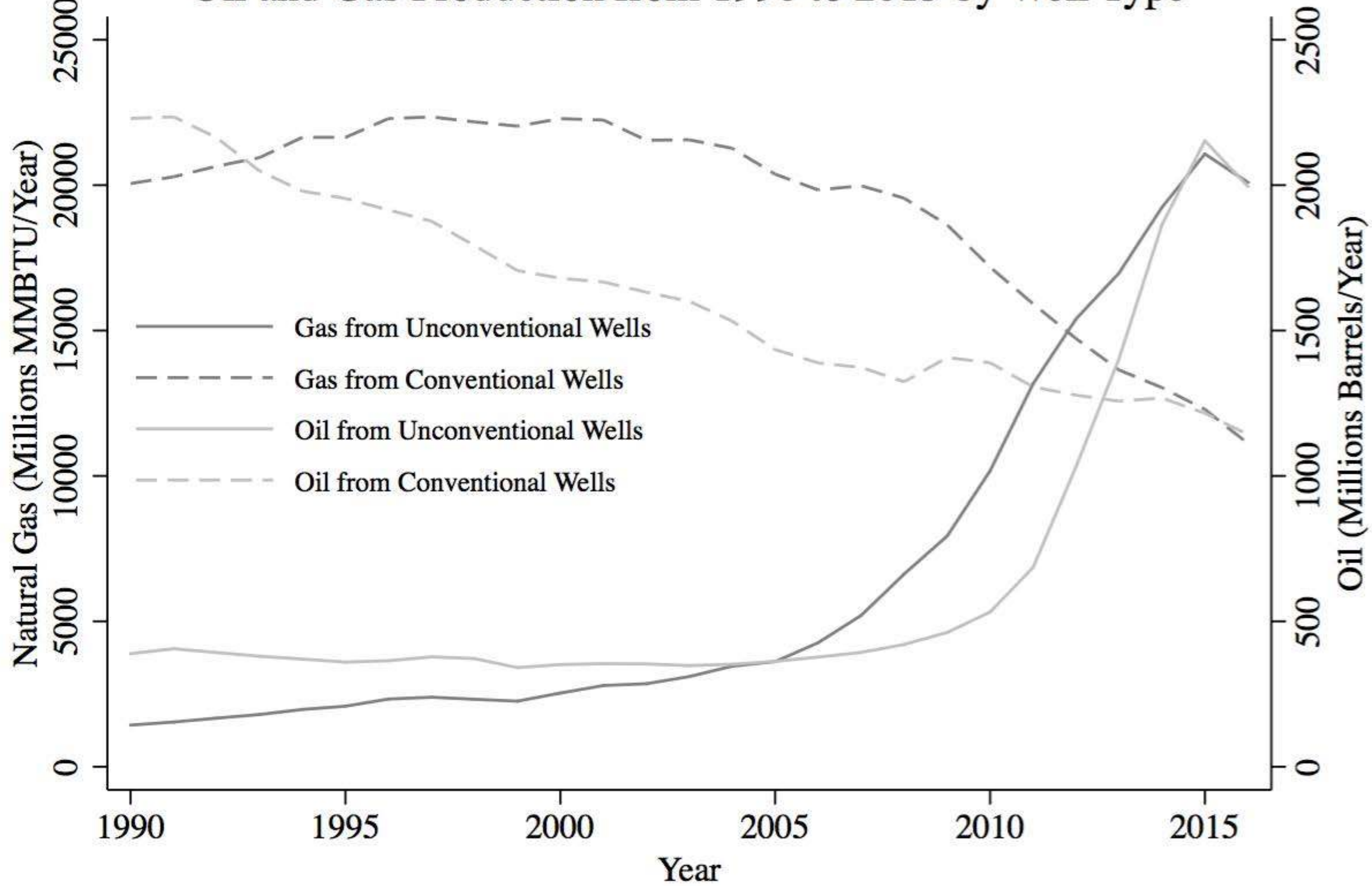
Figure 1 -
Predicted Per-Capita Shale Oil and Gas Reserves by Commuting Zone
in States Overlapping with Major Shale Plays (Thousands of MMBTUs per Person)



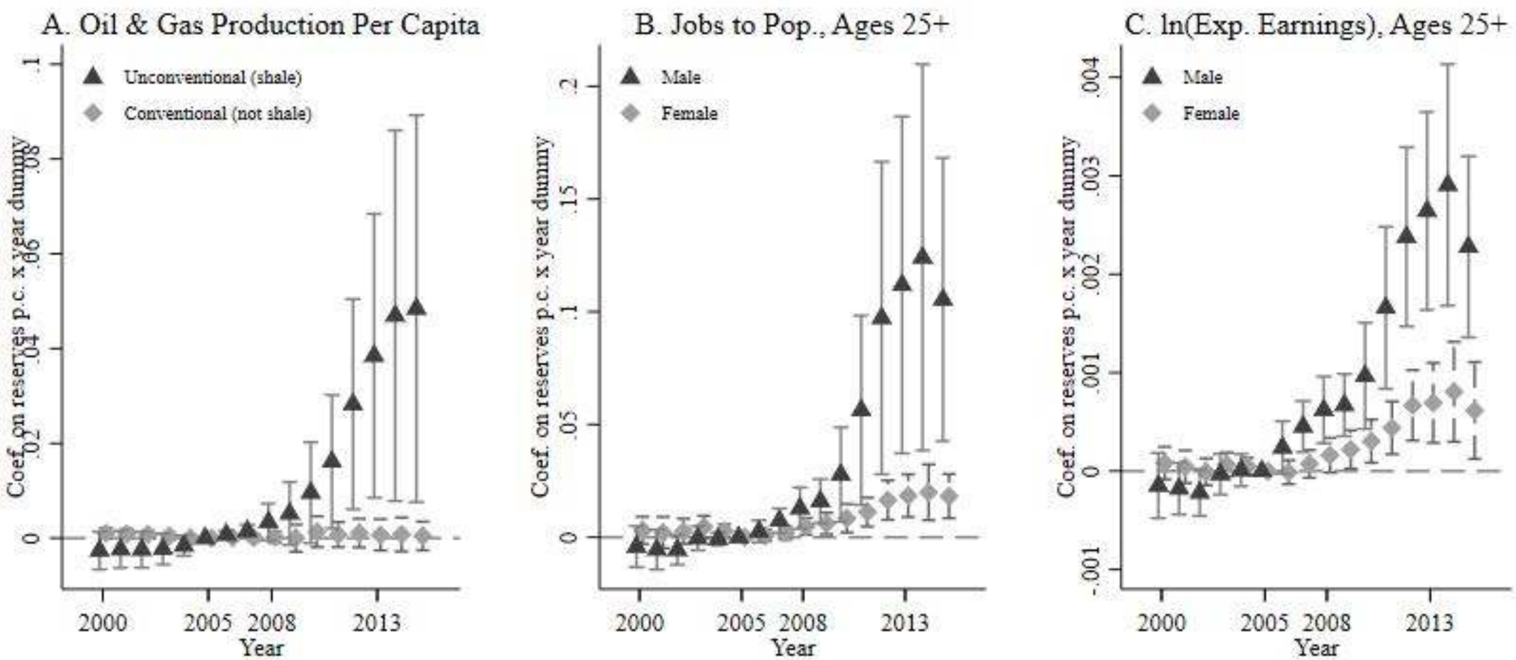
Notes: Figure gives estimated shale oil and gas reserves divided by 2000 population for commuting zones (1990 definition) in 17 states where any CZ lies atop a major shale play. Estimates of shale oil and gas reserves were derived by overlaying recent (2011) EIA maps of shale plays to commuting zones, separately for oil and gas, and allocating maximum EIA estimates of play reserves to commuting zones based on the fraction of each play that they contain. See Data Appendix.

Figure 2 -
Oil and Gas Production from 1990 to 2015 by Well Type



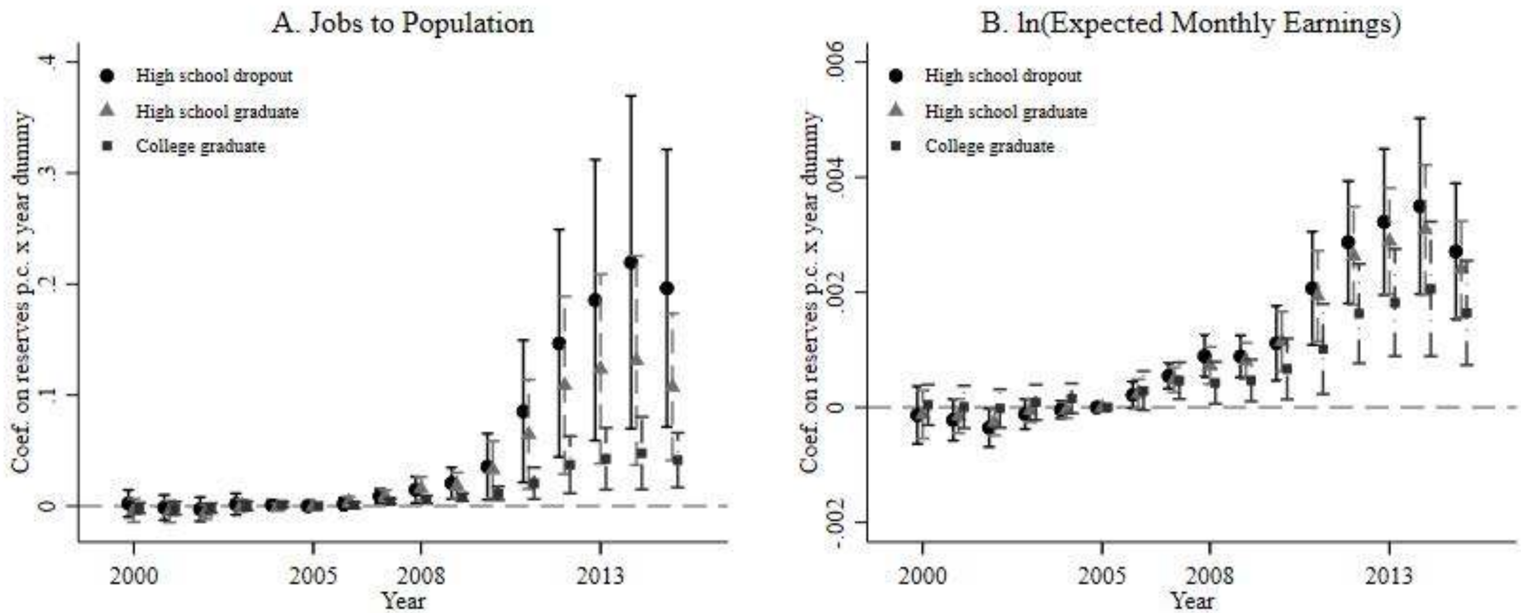
Notes: Data from DrillingInfo.com. Unconventional production is defined as production from horizontal and directional wells. Conventional production is defined as production from vertical and unknown wells. Sample consists of all 48 contiguous states. See Data Appendix.

Figure 3 -
Impacts of Fracking on Oil and Gas Production and the Local Labor Market



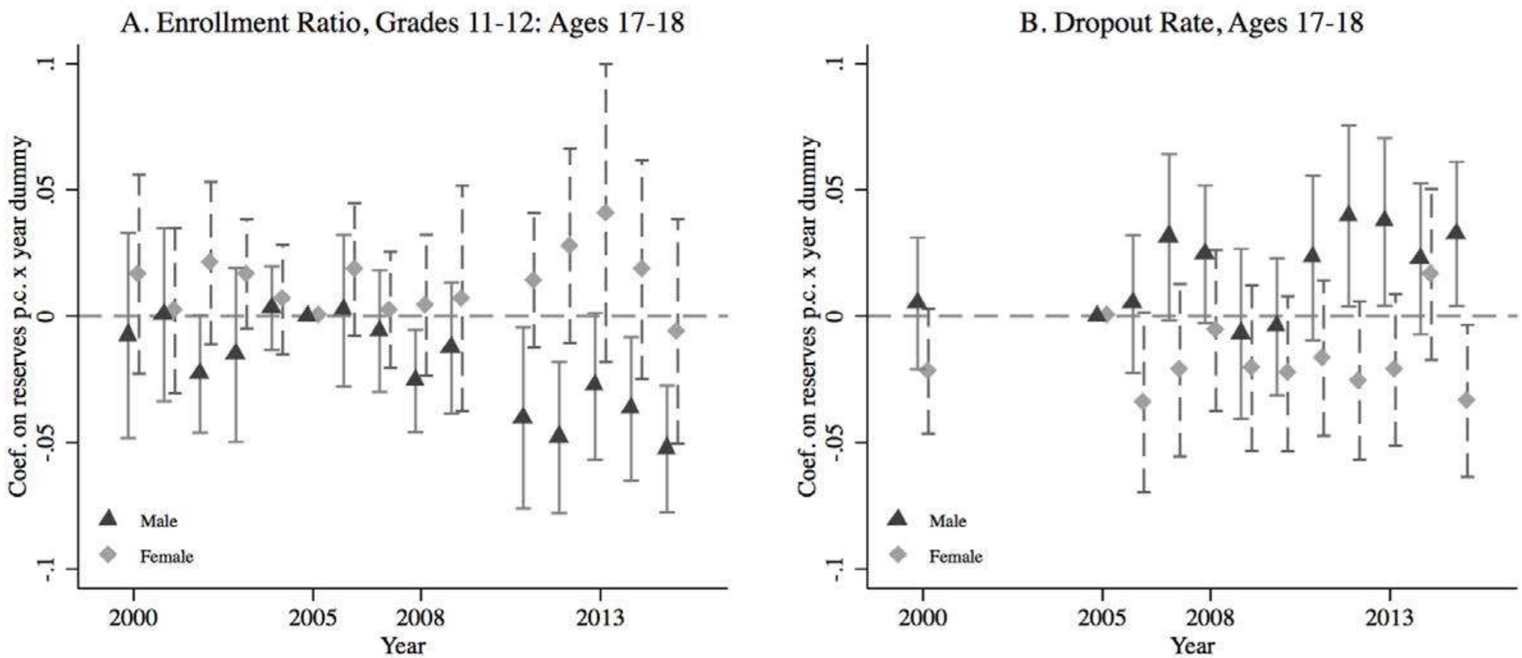
Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that also include state-by-year fixed effects and commuting zone fixed effects. Each commuting zone is given equal weight in the estimation, and inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from DrillingInfo (Panel A) and the Quarterly Workforce Indicators (Panels B and C) and span 2000-2015; see Data Appendix. Sample is limited to 202 CZs in the 14 analysis states.

Figure 4 -
The Effect of Fracking on Adult Male Jobs and Expected Earnings, by Educational Attainment



Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that also include state-by-year fixed effects and commuting zone fixed effects. Each commuting zone is given equal weight in the estimation, and inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from the Quarterly Workforce Indicators (Panel B) and span 2000-2015; see Data Appendix. Sample is limited to 202 CZs in the 14 analysis states.

Figure 5 -
The Effect of Fracking on High School Enrollment & Dropout



Notes: Graph plots coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 population) from regressions that also include state-by-year fixed effects, commuting zone fixed effects, and interactions between year dummies and each of the year 2000 CZ characteristics summarized in Table 1 Panel D. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. In Panel B, regressions are weighted by the number of Census or ACS respondents used to generate the commuting zone-by-year mean dropout rates. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data in Panel A are from the Common Core of Data from 1999-00 to 2014-15, and enrollment ratio is defined as the ratio of 11th and 12th grade enrollment to the 17-18-year-old population. Data in Panel B are from the 2000 Census and the 2005-2015 ACS PUMS, and dropout is defined as being not currently enrolled and without a high school degree. See Data Appendix. Sample is limited to 200 CZs (Panel A) and 202 CZs (Panel B) in the 14 analysis states.

Table 1 -
Year 2000 Commuting Zone Characteristics and their
Association with the Presence and Intensity of Shale Oil and Gas Reserves

	Mean (sd) (1)	Coef (se) on *reserves* * (2)	Coef x mean *res* | *res* >0 (*% of mean*) (3)
	A. Oil and Gas Industry Characteristics		
Shale oil & gas reserves per capita (*reserves*)*	18.53	-	-
	(52.34)		
reserves | *reserves* >0	42.06	-	-
	(72.50)		
Shale oil & gas production per capita*	0.231	-0.0002	-0.0084
	(0.890)	(0.0008)	*(-3.6)*
Conventional oil & gas production per capita*	0.961	0.0005	0.0210
	(3.160)	(0.0020)	*(2.2)*
	B. Employment Characteristics (x100)		
Jobs to population, male ages 25+	55.8	-0.0385	-1.62
	(11.4)	(0.0120)	*(-2.9)*
Jobs to population, female ages 25+	49.5	-0.0217	-0.91
	(8.1)	(0.0074)	*(-1.8)*
ln(expected monthly earnings, male ages 25+)	7.8	-0.0017	-0.07
	(0.3)	(0.0005)	
ln(expected monthly earnings, female ages 25+)	7.1	-0.0009	-0.04
	(0.3)	(0.0003)	
	C. Educational Characteristics (x100)		
Enrollment to population, male 11+12: 17+18	75.4	0.0089	0.37
	(11.7)	(0.0209)	*(0.5)*
Enrollment to population, female 11+12: 17+18	80.4	-0.0023	-0.10
	(12.8)	(0.0214)	*(-0.1)*
High school dropout, male ages 17-18 (%)	10.45	-0.0062	-0.261
	(3.42)	(0.0097)	*(-2.5)*
High school dropout, female ages 17-18 (%)	8.22	-0.0086	-0.362
	(2.94)	(0.0071)	*(-4.4)*

Table 1 (cont'd) -
Baseline Commuting Zone Characteristics and their
Association with the Presence and Intensity of Shale Oil and Gas Reserves

	Mean (sd) (1)	Coef (se) on *reserves* * (2)	Coef x mean *res* | *res* >0 (*% of mean*) (3)
	D. Other CZ Characteristics		
Black population share (%)	5.23	-0.0081	-0.341
	(7.61)	(0.0045)	*(-6.5)*
Hispanic population share (%)	12.80	0.0598	2.515
	(19.09)	(0.0352)	*(19.6)*
Unemployment rate (%)	3.50	0.0032	0.135
	(1.27)	(0.0035)	*(3.8)*
Child poverty rate (%)	19.00	0.0274	1.153
	(7.33)	(0.0161)	*(6.1)*
Median annual household income ($2012)	32,318	-21.0541	-886
	(5,744)	(6.8168)	*(-2.7)*
Child disability rate (%)	7.63	-0.0047	-0.198
	(1.44)	(0.0017)	*(-2.6)*
Single parent population share (%)	11.24	0.0053	0.223
	(2.56)	(0.0064)	*(2.0)*
English-speaking only share (%)	86.35	-0.0490	-2.061
	(15.84)	(0.0324)	*(-2.4)*
Population density	99	-0.1467	-6.2
	(177)	(0.0494)	*(-6.2)*
N (Commuting Zones)	202	202	202

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the Quarterly Workforce Indicators (QWI) for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for all variables except the enrollment-to-population ratio, where there are 200 CZs). The unit of observation is CZ-year (Panels A and D) or CZ-year-sex (Panels B and C). Data were drawn from DrillingInfo (Panel A oil and gas production), the QWI (Panel B jobs numbers), the Common Core of Data (Panel C enrollment numbers), the 2000 Census PUMS (Panel C dropout rates), and 2000 Census county-level tabulations (Panel D and the population figures for the jobs and enrollment to population ratios). Column 2 regressions include state fixed effects, and standard errors are heteroskedasticity-robust. * Measured in 1000s of MMBTUs.

Table 2 -
The Effect of Shale Oil and Gas Reserves and
the Introduction of Fracking on the Jobs to Population Ratio, by Sex

Dependent Variable:	Jobs to Population Ratio Ages 25 and over (x100)		
	Male	Female	Male-Female Difference
	(1)	(2)	(3)
	A. Jobs to Population		
Average effect 2011-15, *reserves*>0	4.28	0.60	3.68
	(1.77)	(0.28)	(1.57)
Coefficient on:			
Shale reserves per capita x 2006-10	0.0161	0.0019	0.0142
	(0.0078)	(0.0038)	(0.0071)
Shale reserves per capita x 2011-15	0.1017	0.0143	0.0874
	(0.0421)	(0.0067)	(0.0373)
p : = coefs. (across yr. groups)	*0.014*	*0.003*	*0.020*
p : = coefs. (across sex, 2011-15)			*0.020*
Observations	3,232	3,232	6,464
R-squared	0.877	0.957	0.898
	B. Ln(Expected Monthly Earnings)		
Average effect 2011-15, *reserves*>0	0.104	0.026	0.078
	(0.027)	(0.010)	(0.018)
Coefficient on:			
Shale reserves per capita x 2006-10	0.0007	0.0001	0.0006
	(0.0002)	(0.0001)	(0.0002)
Shale reserves per capita x 2011-15	0.0025	0.0006	0.0019
	(0.0006)	(0.0002)	(0.0004)
p : = coefs. (across yr. groups)	*0.000*	*0.002*	*0.000*
p : = coefs. (across sex, 2011-15)			*0.000*
Observations	3,232	3,232	6,464
R-squared	0.936	0.954	0.960

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data on jobs and monthly earnings for individuals ages 25 and over are from the 2000-2015 QWI and correspond to unweighted averages of beginning of quarter figures reported throughout the year, and data on the population ages 25 and over are from the Surveillance, Epidemiology, and End Results (SEER) program. Expected monthly earnings multiply reported monthly earnings by the jobs-to-population ratio. All regressions include state-by-year fixed effects, CZ fixed effects, and time-varying effects of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Regressions in column 3 also include interactions between these variables and a dummy for male; coefficients on the interactions with the male dummy are given in the table. Standard errors (in parentheses) are clustered on CZ.

Table 3 -
The Effect of Shale Oil and Gas Reserves and the Introduction of Fracking on Jobs to Population and Expected Monthly Earnings, by Sex and Educational Attainment

		Average effect 2011-15, *reserves* >0			
	Overall	High School Dropouts	High School Graduates	College Attendees	College Graduates
Dependent variable:	(1)	(2)	(3)	(4)	(5)
			A. Men (N=3,232)		
Jobs to Population (x100)	4.28	7.01	4.61	6.06	1.62
	(1.77)	(2.93)	(1.98)	(2.48)	(0.67)
p : = coefs (v. dropout)			*0.063*	*0.357*	*0.021*
ln(Expected monthly earnings)	0.104	0.127	0.113	0.106	0.067
	(0.027)	(0.034)	(0.025)	(0.025)	(0.024)
p : = coefs (v. dropout)			*0.185*	*0.060*	*0.000*
			B. Women (N=3,232)		
Jobs to Population (x100)	0.60	0.61	0.55	1.33	0.15
	(0.28)	(0.75)	(0.29)	(0.34)	(0.18)
p : = coefs (v. dropout)			*0.910*	*0.220*	*0.520*
ln(Expected monthly earnings)	0.026	0.051	0.031	0.029	0.009
	(0.010)	(0.019)	(0.012)	(0.009)	(0.009)
p : = coefs (v. dropout)			*0.041*	*0.075*	*0.003*
			C. Male-Female Difference (N=6,464)		
Jobs to Population (x100)	3.68	6.40	4.06	4.73	1.47
	(1.57)	(2.80)	(1.77)	(2.24)	(0.54)
p : = coefs (v. dropout)			*0.131*	*0.227*	*0.040*
ln(Expected monthly earnings)	0.078	0.076	0.083	0.076	0.058
	(0.018)	(0.024)	(0.016)	(0.017)	(0.017)
p : = coefs (v. dropout)			*0.531*	*0.948*	*0.198*

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data on jobs and monthly earnings for individuals ages 25 and over are from the 2000-2015 QWI and correspond to unweighted averages of beginning of quarter figures reported throughout the year; data on the population ages 25 and over are from SEER; and estimates of education shares in the poulation ages 25-64 are from the Census and ACS. Expected monthly earnings multiply reported monthly earnings by the group-specific jobs-to-population ratio. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 (Panels A and B) or a fully-interacted version of model 2 using pooled data (Panel C). All regressions include state-by-year and CZ fixed effects and time-varying effects of the CZ observables summarized in Table 1 Panel D. Each CZ is given equal weight in the estimation. Standard errors clustered on CZ are in parentheses.

Table 4 -
The Effect of Shale Oil and Gas Reserves and the Introduction of Fracking on 11th and 12th Grade Enrollment and High School Dropout Rates of 17- and 18-Year-Olds

	Male	Female	Male-Female Difference
	(1)	(2)	(3)
	A. Enrollment Ratio, 11th & 12th: Ages 17-18		
Average effect 2011-15, *reserves*>0	-1.43	0.36	-1.79
	(0.63)	(0.81)	(1.09)
Coefficient on:			
Shale reserves per capita	-0.0035	-0.0024	-0.0010
x 2006-10	(0.0163)	(0.0166)	(0.0122)
Shale reserves per capita	-0.0340	0.0085	-0.0425
x 2011-15	(0.0151)	(0.0192)	(0.0259)
p : = coefs. (across yr. groups)	*0.07*	*0.39*	*0.08*
p : = coefs. (across sex, 2011-15)			*0.10*
Observations	3,000	3,000	6,000
R-squared	0.764	0.777	0.785
	B. Dropout Rate, Ages 17-18		
Average effect 2011-15, *reserves*>0	1.12	0.12	1.00
	(0.36)	(0.32)	(0.45)
Coefficient on:			
Shale reserves per capita	0.0054	-0.0022	0.0076
x 2006-10	(0.0082)	(0.0072)	(0.0098)
Shale reserves per capita	0.0267	0.0030	0.0238
x 2011-15	(0.0086)	(0.0075)	(0.0107)
p : = coefs. (across yr. groups)	*0.01*	*0.50*	*0.13*
p : = coefs. (across sex, 2011-15)			*0.03*
Observations	2,424	2,424	4,848
R-squared	0.492	0.437	0.484

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 out of 259 CZs for the dropout rate and 200 of 259 CZs for the enrollment-to-population ratio). The unit of observation is CZ-year-sex. Enrollment data in Panel A are from the 1999-00 to 2014-15 Common Core of Data (missing 2009-10); enrollment figures are divided by SEER-based estimates of the CZ's 17- and 18-year-old population. Dropout data are from the 2000 Census PUMS and the 2005 through 2015 ACS PUMS. All regressions include state-by-year and CZ fixed effects, as well as time-varying effects of the 2000 Census CZ characteristics listed in Table 1 Panel D. Regressions in column 3 also include interactions between these variables and a dummy for male; coefficients on the interactions with the male dummy are given in the table. Panel B regressions are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean dropout rates. Standard errors (in parentheses) are clustered on CZ.

Table 5 -
Sensitivity of the Estimates for Dropout, Enrollment, and Earnings to Choice of Specification

	Average effect in 2011-15, *reserves*>0					
Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
			A. Baseline Model			
Preferred specification	-1.43	0.36	1.12	0.12	0.104	0.026
	(0.63)	(0.81)	(0.36)	(0.32)	(0.027)	(0.010)
p : = coefs. (across sex, 2011-15)	*0.10*		*0.03*		*0.000*	
			B. Dropping Controls and Changing Weighting			
No time-varying effects of CZ observables	-1.25	0.37	1.13	0.35	0.109	0.030
	(0.63)	(0.62)	(0.43)	(0.37)	(0.024)	(0.009)
p : = coefs. (across sex, 2011-15)	*0.15*		*0.10*		*0.000*	
Year fixed effects	-1.49	0.01	1.47	0.57	0.117	0.032
	(0.62)	(0.80)	(0.40)	(0.32)	(0.031)	(0.013)
p : = coefs. (across sex, 2011-15)	*0.19*		*0.03*		*0.000*	
Unweighted (changes dropout only)	-1.43	0.36	1.03	-0.06	0.104	0.026
	(0.63)	(0.81)	(0.23)	(0.22)	(0.027)	(0.010)
	0.10		*0.00*		*0.000*	
			C. Variation in Timing			
Use first frack date for largest play in state	-1.00	0.22	0.68	0.44	0.075	0.019
	(0.46)	(0.68)	(0.30)	(0.29)	(0.020)	(0.008)
p : = coefs. (across sex, post-fracking)	*0.16*		*0.49*		*0.000*	
Long difference: 2000 v. 2011-15	-1.41	0.10	1.04	0.24	0.105	0.025
	(1.11)	(0.67)	(0.41)	(0.35)	(0.028)	(0.011)
p : = coefs. (across sex, 2011-15)	*0.17*		*0.12*		*0.000*	

Table 5 (cont'd) -
Sensitivity of the Estimates for Dropout, Enrollment, and Earnings to Choice of Specification

	Average effect in 2011-15, *reserves*>0					
Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
			D. Alternative Reserve Estimates			
Value of reserves using 2003 prices	-1.40	0.37	1.12	0.14	0.102	0.025
(max gas price:oil price, 1000s of $2012)	(0.62)	(0.80)	(0.36)	(0.31)	(0.027)	(0.011)
p : = coefs. (across sex, 2011-15)	*0.10*		*0.03*		*0.000*	
Value of reserves using 2012 prices	-1.05	0.38	0.86	0.27	0.080	0.021
(max oil price:gas price, 1000s of $2012)	(0.49)	(0.63)	(0.28)	(0.22)	(0.026)	(0.010)
p : = coefs. (across sex, 2011-15)	*0.09*		*0.06*		*0.001*	
Simulated reserves in 2008	-0.66	0.44	0.58	0.17	0.070	0.019
(1000s of MMBTUs)	(0.59)	(0.62)	(0.27)	(0.22)	(0.022)	(0.009)
p : = coefs. (across sex, 2011-15)	*0.25*		*0.26*		*0.001*	

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for each outcome except the enrollment-to-population ratio, where there are 200 CZs). The unit of observation is CZ-year-sex. Unless otherwise noted (Panel B), all regressions include state-by-year and CZ fixed effects, as well as time-varying effectsd of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 as described in the text. Regressions for dropout rates (columns 3 and 4) are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean dropout rates. Standard errors (in parentheses) are clustered on CZ.

Table 6 -
Education, Location, or Demographics?

	Effect in 2011-15 in average CZ with *reserves* >0		*p* : =coefs, across
	Male	Female	sex, 2011-15
	(1)	(2)	(3)
		A. Baseline Model	
Preferred specification (dropout)	1.12	0.12	*0.03*
	(0.36)	(0.32)	
		B. Effects of Migration	
Outcome is dropout (ages 17-18) *and* recent migrant	0.35	0.04	*0.34*
	(0.24)	(0.21)	
Outcome is dropout (ages 17-18) *and not* recent migrant	0.77	0.08	*0.04*
	(0.25)	(0.25)	
		C. Additional Controls	
Dropout: Control for shares black, Hispanic, and recent migrant in CZ population	1.13	0.14	*0.03*
	(0.36)	(0.32)	
Dropout: Control for state CSL of 17 or 18 (time-varying) x *reserves*	1.20	0.07	*0.02*
	(0.38)	(0.33)	
Enrollment ratio: Control for state CSL of 17 or 18 (time-varying) x *reserves*	-1.68	-0.39	*0.29*
	(0.83)	(0.64)	

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 out of 259 CZs for the dropout rate and 200 of 259 CZs for the enrollment-to-population ratio). The unit of observation is CZ-year-sex. All regressions include state-by-year and CZ fixed effects, as well as time-varying effects of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 as described in the text. Regressions where dropout rates are the dependent variable are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex means. Standard errors (in parentheses) are clustered on CZ.

Table 7 -
Alternative Mechanisms: Job Opportunities for Teenagers?

	Year 2000	Model (2) Estimate for *reserves* x		*p* : =coefs			Average Effect in
	Mean	2006-10	2011-15	x yr. grp.	x sex	v. 14-18	2011-15
Dependent variable:	(1)	(2)	(3)	(4)	(5)		(6)
		A. Jobs to Population, by Sex and Age					
Males Ages 14-18	19.1	0.0057	0.0074	*0.67*			0.31
		(0.0026)	(0.0045)				(0.19)
Females Ages 14-18	22.5	0.0026	-0.0005	*0.47*	*0.03*		-0.02
		(0.0026)	(0.0051)				(0.21)
Males Ages 19-24	53.4	0.0299	0.1140	*0.02*		*0.01*	4.79
		(0.0090)	(0.0422)				(1.77)
Females Ages 19-24	60.0	0.0207	0.0489	*0.00*	*0.09*	*0.00*	2.06
		(0.0077)	(0.0098)				(0.41)
Males Ages 25-34	63.1	0.0225	0.1383	*0.01*		*0.01*	5.82
		(0.0132)	(0.0562)				(2.36)
Females Ages 25-34	63.6	0.0065	0.0218	*0.01*	*0.03*	*0.01*	0.92
		(0.0059)	(0.0088)				(0.37)
		B. ln(Expected Monthly Earnings), by Sex and Age					
Males Ages 14-18	4.71	0.0009	0.0021	*0.01*			0.089
		(0.0003)	(0.0006)				(0.027)
Females Ages 14-18	4.71	0.0005	0.0007	*0.46*	*0.00*		0.030
		(0.0002)	(0.0004)				(0.017)
Males Ages 19-24	6.70	0.0013	0.0034	*0.00*		*0.00*	0.141
		(0.0003)	(0.0006)				(0.027)
Females Ages 19-24	6.49	0.0005	0.0013	*0.00*	*0.00*	*0.07*	0.057
		(0.0001)	(0.0003)				(0.014)
Males Ages 25-34	7.48	0.0008	0.0030	*0.00*		*0.07*	0.126
		(0.0003)	(0.0008)				(0.033)
Females Ages 25-34	7.11	0.0003	0.0008	*0.00*	*0.00*	*0.69*	0.036
		(0.0002)	(0.0003)				(0.012)

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data are from the 2000-2015 QWI and SEER. All regressions include state-by-year and CZ fixed effects, as well as time-varying effects of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Each CZ is given equal weight. Standard errors (in parentheses) are clustered on CZ.

Table 8 -
Alternative Mechanisms: Changes in School Resources?

Dependent variable:	Year 2000 Mean (1)	Model (2) Estimate for *reserves* x 2006-10 (2)	2011-15 (3)	p : =coefs (4)	Average Effect in 2011-15 (5)
Class size	18.59	0.0129	0.0002	*0.32*	0.01
		(0.0129)	(0.0105)		(0.44)
ln(Total spending per pupil)	8.95	0.0001	0.0001	*0.75*	0.006
		(0.0001)	(0.0002)		(0.009)
ln(Total revenue per pupil)	8.95	0.0001	0.0003	*0.46*	0.013
		(0.0002)	(0.0002)		(0.008)
ln(Federal revenue per pupil)	6.44	-0.0000	-0.0004	*0.05*	-0.018
		(0.0001)	(0.0002)		(0.010)
ln(State revenue per pupil)	8.21	0.0001	-0.0006	*0.01*	-0.023
		(0.0003)	(0.0003)		(0.014)
ln(Local revenue per pupil)	7.97	0.0001	0.0015	*0.02*	0.063
		(0.0002)	(0.0007)		(0.028)

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data were drawn from the Common Core of Data for the academic years 1999-2000 to 2014-2015. All regressions include state-by-year and CZ fixed effects, as well as time-varying effects of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Each CZ is given equal weight. Standard errors (in parentheses) are clustered on CZ.

Data Appendix

A. Estimating Shale Gas and Oil Reserves

Our preferred approach to predicting CZ reserves relies on the maximum ever (across 2008 to 2015 for gas and 2011 to 2015 for oil) reported reserves for each major shale play, published by the Energy Information Administration (EIA), and the 2011 EIA maps of shale plays.[1] We overlay these maps to counties, separately for oil and gas, and allocate maximum reserves to counties based on the fraction of each play that they contain, following a process similar to that of Maniloff and Mastromonaco (2017), who study the local economic impacts of fracking.[2] A given CZ's oil (gas) reserves are then the sum of these prorated oil (gas) reserves across all counties the CZ contains.[3] To combine oil and gas reserves, we convert these predicted reserves into millions of British Thermal Units (MMBTUs), which capture the amount of heating energy that they contain.[4] Finally, we normalize each CZ's combined predicted oil and gas reserves by its 2000 population to arrive at predicted shale reserves per capita.

B. Data on Gas and Oil Production

Figure 2, Figure 3A, and Appendix Table A1 were estimated using CZ-by-year aggregates of monthly well-level production data licensed to us under a special agreement with DrillingInfo. The aggregate data give production (thousands of cubic feet of natural gas or barrels of oil) by well type (horizontal or directional versus vertical or unknown).

Following prior research (Feyrer, Mansur, and Sacerdote, 2017), we classify production from horizontal and directional (unconventional) wells as fracking, or as coming from shale. As with our reserve measure, we combine oil and gas production using the conversion to MMBTUs. In Appendix Table A1, however, we also consider the value of production (converted to real 2012 dollars using the energy CPI) as an outcome. To reduce imprecision in the estimates arising from some extreme outliers in these data, we follow Maniloff and Mastromonaco (2017) in also taking the inverse hyperbolic sine (IHS) of these measures. The IHS effectively allows us to take the natural logarithm of production – and coefficients can be interpreted much as they would be in a log-linear model as production levels grow – but retain observations with zero production in the estimation sample.

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[1] Reserves by play for, for example, 2013, are reported at http://www.eia.gov/naturalgas/crudeoilreserves/pdf/table_2.pdf and http://www.eia.gov/naturalgas/crudeoilreserves/pdf/table_4.pdf. The shapefiles for play boundaries can be found at http://www.eia.gov/pub/oil_gas/natural_gas/analysis/publications/maps/shalegasplay.zip. We test the robustness of our results to changing various aspects of the prediction in Table 5.

[2] Thus, we assign a county *x*% of a shale play's estimated reserves if it accounts for *x*% of its land area. Unlike Maniloff and Mastromonaco (2017), we use more frequent data and CZs rather counties. We also use the estimated reserves rather than the fraction of each CZ with reserves to better capture fracking potential. For example, two CZs with very different reserves would look identical under the latter measure. The latter measure also cannot easily accommodate the fact that some labor markets lie atop multiple shale plays.

[3] We allocate counties to CZs (1990 boundaries) using the crosswalk provided by David Dorn: http://www.ddorn.net/data/cw_cty_czone.zip.

[4] We use the production conversion factors reported for 2012 by the EIA (http://www.eia.gov/forecasts/aeo/pdf/appg.pdf): 1,022 BTUs per cubic foot of gas and 5.85 MMBTUs per barrel of oil.

C. Census and ACS Microdata

Overview and Key Variables

The 2000 Census and the 2005 through 2015 American Community Survey (ACS) Public Use Microdata Samples (PUMS) (Ruggles et al., 2015) provide individual-level data on gender, age, school enrollment, and educational attainment, and identify local geography down to the Public Use Microdata Area (PUMA) level. PUMAs are not the same as CZs, but can be allocated to CZs based on the division of county population across PUMAs and the mapping between CZs and counties.[5] Thus, we are able to estimate sex-specific dropout rates of teenagers at the CZ level for two pre-fracking years (2000 and 2005) and ten consecutive post-fracking years (2006 through 2015). Unfortunately, no information on PUMA is provided in the public-use ACS PUMS for 2001 to 2004.[6]

We define dropouts as those who have not recently been enrolled in school and do not have a high school degree and limit attention to 17- to 18-year-olds in our main analysis. We focus on the population that is of high school age to mitigate bias from selective migration. That is, by including 19-year-olds or older adults, we would be more concerned that what might appear to be an impact of fracking on education decisions is really an impact of fracking on location choices of existing dropouts. We are interested in how fracking has affected the level of dropout in the economy, not how it has affected the geographic distribution of existing dropouts. Fortunately, by using microdata, we are able to explore the influence of migration on our estimates directly, which we do in Table 6.

Considerations

The reporting period for school enrollment is the prior three months in the ACS (which is conducted in all months of the year) and February 1 in the Census (which is conducted on April 1). These differences in generate several challenges in comparing dropout rates of 17- and 18-year-olds across the two surveys, since survey month is not publicly reported in the ACS. First, because school years generally last 180 days, starting in September and ending in May, there is some potential that we misclassify ACS respondents interviewed in the late summer as dropouts, since we do not observe the interview date in the public-use data. Because our analysis includes state-by-year fixed effects, it will account for any resulting bias provided that it does not vary within states over time. Second, with a new school year starting in September, a higher share of the 17- and 18-year-olds surveyed in the fall – and therefore a higher share of 17- and 18-year-olds in the ACS – will be of age to be enrolled in 12th grade or beyond. If teens sometimes make dropout decisions based on grade rather than age, and if that tendency happens to be correlated with *reserves*, our estimates from the Census/ACS analysis could be biased.

D. Common Core of Data

[5] Crosswalks between 1990 and 2000 PUMAs and CZs (relevant for the 1990 Census, 2000 Census, and the 2005-2011 ACS) are available on David Dorn's website: http://www.ddorn.net/data/cw_puma2000_czone.zip. We create an analogous crosswalk between 2010 PUMAs and CZs (relevant for the 2012-2015 ACS) using data on the division of 2010 county population across 2010 PUMAs from the Missouri Census Data Center (http://mcdc.missouri.edu) and the county-CZ crosswalk (http://www.ddorn.net/data/cw_cty_czone.zip).

[6] PUMA codes are not reported in the public-use ACS files for these years.

The Common Core of Data (CCD) is the primary federal database on public schools in the United States. We used the Department of Education's "ELSi tableGenerator"[7] to gather aggregated annual county-level CCD data on the number of students by grade and gender (for calculation of the enrollment-to-population ratio)[8] and the total number of students and teachers (for calculation of the student-teacher ratio/class size).

For several states and years, we are missing sex-specific enrollment data. In these cases, we impute missing values. We first attempt to impute with the product of total enrollment and the sex-specific share in the enrolled population (averaged over 2000-2016). However, this only works when total enrollment is reported. For years where total enrollment is not reported, we impute with linearly interpolated values. These imputations affect a small number of observations.[9] For the purposes of the enrollment-to-population ratio analysis, we also drop two CZs in Nebraska that are large outliers on this variable.

E. QWI Data Description

Overview

The Quarterly Workforce Indicators (QWI) are a set of quarterly labor market indicators made available by the Longitudinal Employer-Household Dynamics (LEHD) at the U.S. Census Bureau. The LEHD covers over 95 percent of U.S. private sector jobs and is compiled from a variety of sources including state Unemployment Insurance (UI) records, the Quarterly Census of Employment and Wages (QCEW), the Census, Social Security Administrative records, and individual tax returns. Linkages across these data sources enable the LEHD to publish QWI data with detailed firm characteristics and worker demographic information for geographies as small as counties. The near-universe and demographic-specific nature of the QWI is particularly useful for our study, which requires data by sex and education level for local labor markets that often have low populations.

The QWI offers advantages over alternative sources of local economic indicators from the ACS and QCEW. The ACS contains rich information on the characteristics of workers, but it is drawn from a one percent sample of the population. Self-reported earnings in low-population areas further disaggregated by educational attainment and sex can therefore be noisy. In addition, the ACS was only consistently administered from 2005 onwards, while the QWI offer data for earlier years, enabling a more robust analysis of labor market outcomes prior to the onset of fracking. The QCEW provides near-universe data spanning the duration of our study, but it does not contain any demographic details that are critical for our study.

Considerations

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[7] Go to https://nces.ed.gov/ccd/elsi/tableGenerator.aspx.

[8] The age- and sex-specific population estimates come from SEER (Surveillance, Epidemiology, and End Results Program): https://seer.cancer.gov/popdata/download.html.

[9] The first imputation affects approximately 90% of enrollment in the average Virginia CZ in 2000, 90% of enrollment in the average Pennsylvania CZ over 2000-04, 8.6% of enrollment in the average New York CZ over 2000-04, and 2.4% of enrollment in the average Ohio CZ over 2000-04. The second imputation is relevant for less than 0.3% of enrollment in the average CZ in all years except 2012 (4%) and 2015 (7%).

The QWI data do have some limitations which are important to consider in the context of our analysis. First, because much of the data is collected via a unique federal-state data sharing collaboration, the years of data availability differ by state. While 14 states in our study have data available for our entire analysis timeframe (2000-2015), three states (Arkansas, Kentucky, and Wyoming) are missing some years of data, particularly at the beginning of the study window. We exclude these states in our analysis, but estimates that include data from all 17 states where available look similar to those reported in the paper. The table below details the years of data available in the QWI for the 17 states where at least one CZ overlaps with a major shale play.

QWI Data Availability

State	Start Date	State	Start Date
AR	2002: Q3	NY	2000: Q1
CO	1993: Q2	OH	2000: Q1
KY	2001: Q1	OK	2000: Q1
LA	1995: Q1	PA	1997: Q1
MD	1999: Q1	TX	1995: Q1
MT	1993: Q1	VA	1998: Q3
ND	1998: Q1	WV	1997: Q1
NE	1999: Q1	WY	2001: Q1
NM	1995: Q3		

Imputations in the published QWI data warrant consideration as well, especially for education-specific data. QWI documentation states that the vast majority of jobs can be matched to personal characteristics for sex and age variables (97 percent) and geographic variables (over 90 percent), so imputations are less of a concern for these variables. On the other hand, education characteristics primarily rely on an imputation model with a statistical match between the Census and LEHD data using a state-specific logit model that contains age categories, earnings categories, and industry dummies for individuals age 14 and older who reported strictly positive wage earnings. Although this method of imputation likely yields estimates that are generally accurate, it may improperly capture dynamic effects specific to the fracking boom. It also adds noise to our education-specific estimates (Table 3).

Finally, the format of the publicly available QWI data limits our analysis in a couple minor ways. First, the online extraction tool for the publicly available QWI data allows for only three separate ways to tabulate data: by sex and age, by sex and education, and by race and ethnicity. We are therefore restricted to analyses based on two-way tabulations. Additionally, data suppression prevents us from undertaking any industry-specific analysis.

Variable Definitions

Note that the QWI is constructed from firm statistics and therefore represents employment and earnings at the job rather than the individual level. This means that individuals who hold multiple jobs will be counted multiple times in the measures described below.

Jobs-to-Population Ratio: To construct the jobs-to-population ratio used in our study, we create a yearly average of the quarterly total employment reported in the QWI and divide that by population estimates[10] corresponding to the sex by age parameters we select in the QWI. The total employment figure used is defined as the count of people employed in a firm at the beginning of the quarter, which is more of a stock measure than the alternative: the count of people employed in a firm any time during the quarter, which is considered more of a flow measure of jobs. Because this measure is not available until the second quarter of a state's reporting, we are able to construct it using only three quarters of data in the first year that data become available for a state (see table). Overall, this choice does not greatly influence our results as the event study figures are similar when using total quarterly jobs to construct our preferred measure.

Earnings: We construct our preferred earnings measure by taking the beginning quarterly payroll metric in the QWI, dividing by beginning payroll employment to obtain an average quarterly earnings measure, and dividing by 3 to translate this measure to a monthly figure. We convert to an annual measure by taking a weighted average across the four quarters, weighting by beginning of quarter employment in a given quarter. Event study estimates do not meaningfully differ when using total payroll earnings versus beginning of quarter earnings.[11]

Data references

Feyrer, James, Erin T. Mansur, and Bruce Sacerdote. 2017. "Geographic Dispersion of Economic Shocks: Evidence from the Fracking Revolution." *American Economic Review* 107(4): 1313-1334.

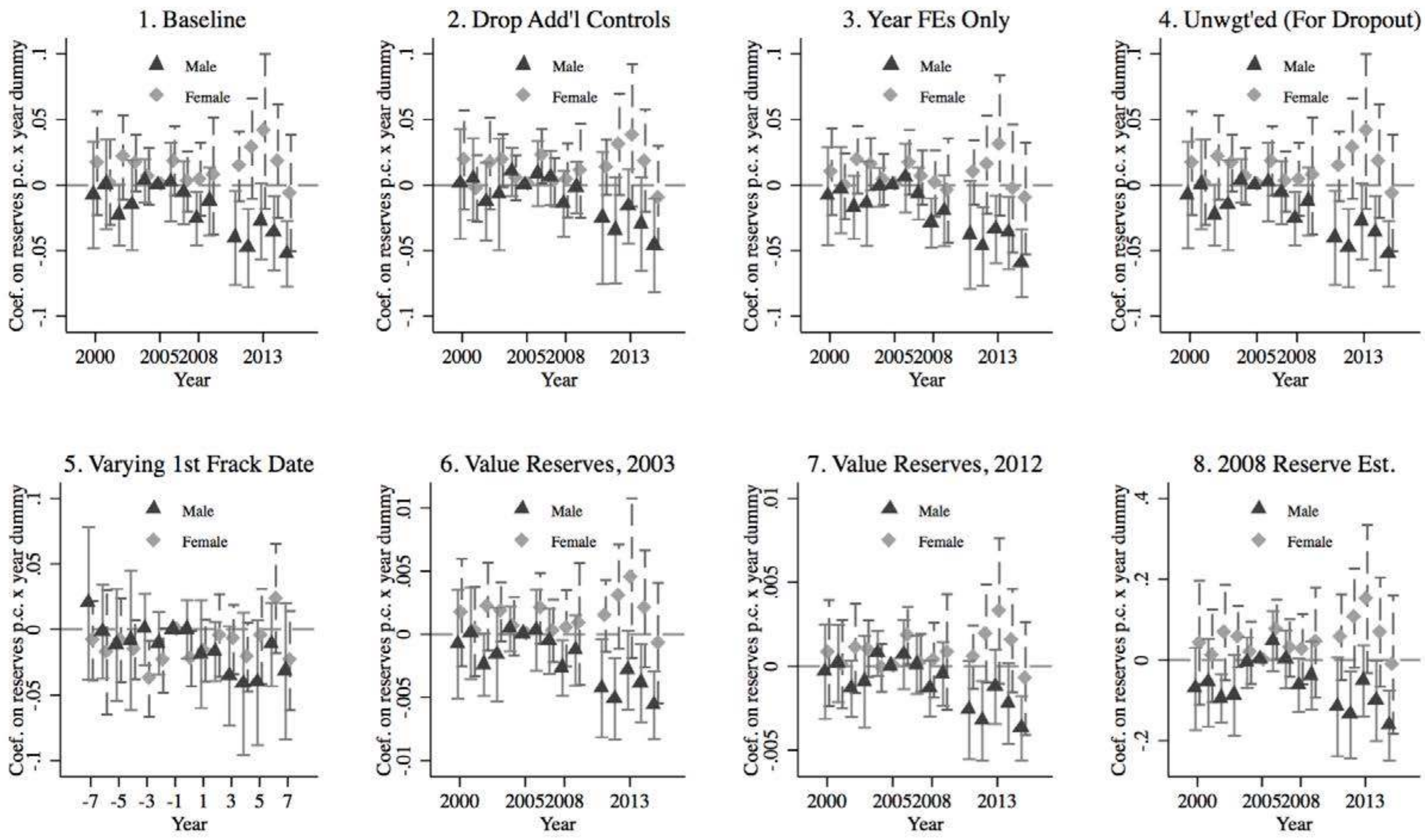
Maniloff, Peter and Ralph Mastromonaco. 2017. "The Local Economic Impacts of Hydraulic Fracturing and Determinants of Dutch Disease." *Resource and Energy Economics* 49: 62-85.

Ruggles, Steven, Katie Genadek, Ronald Goeken, Josiah Grover, and Matthew Sobek. Integrated Public Use Microdata Series: Version 6.0 [dataset]. Minneapolis: University of Minnesota, 2015.

[10] The SEER population estimates are available at https://seer.cancer.gov/popdata/download.html.

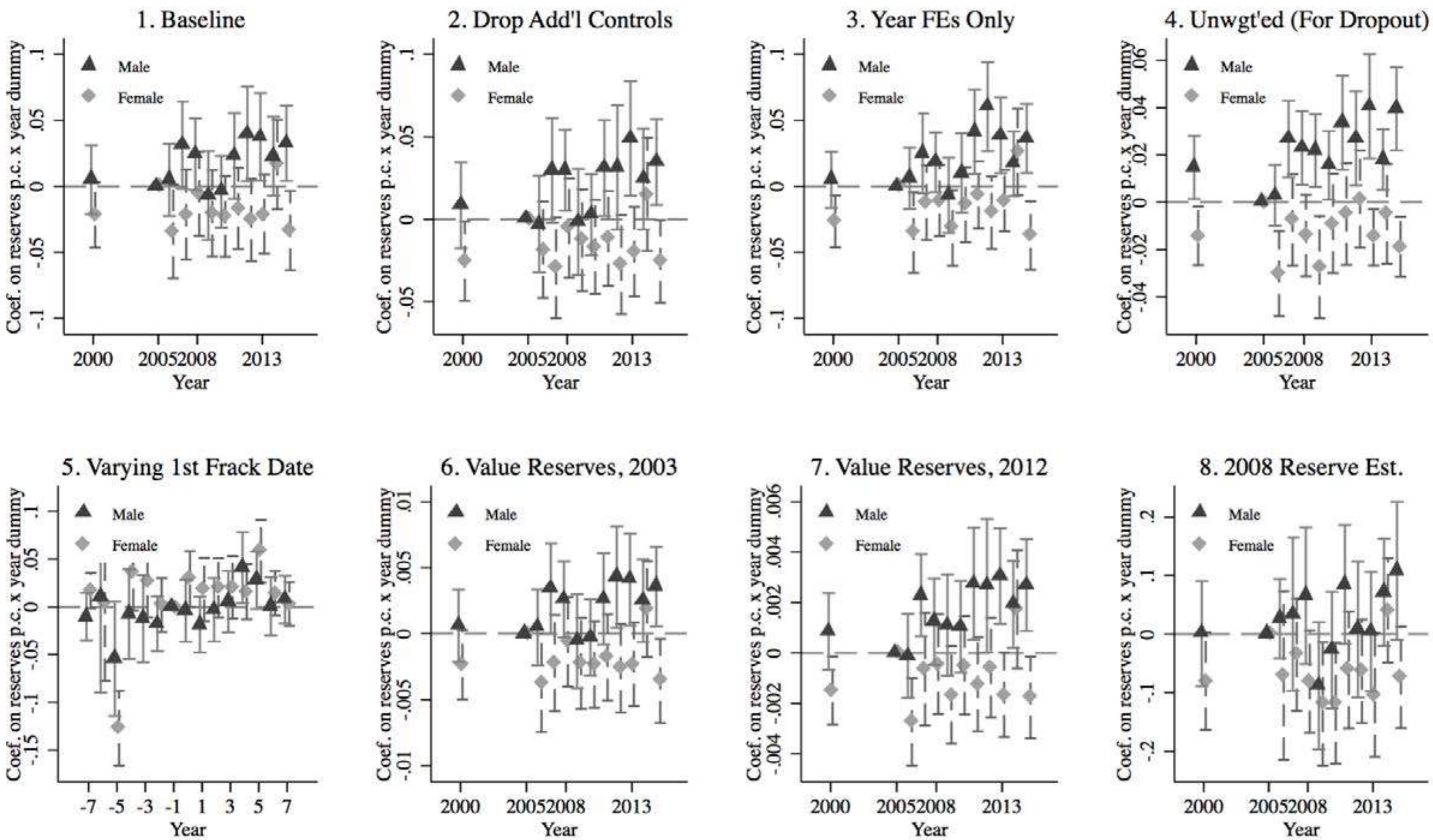
[11] For further explanation of these data, see https://lehd.ces.census.gov/data, https://lehd.ces.census.gov/doc/QWI_101.pdf, https://lehd.ces.census.gov/doc/technical_paper/tp-2006-01.pdf.

Appendix Figure A1 -
Table 5 Specification Checks, Enrollment Ratio



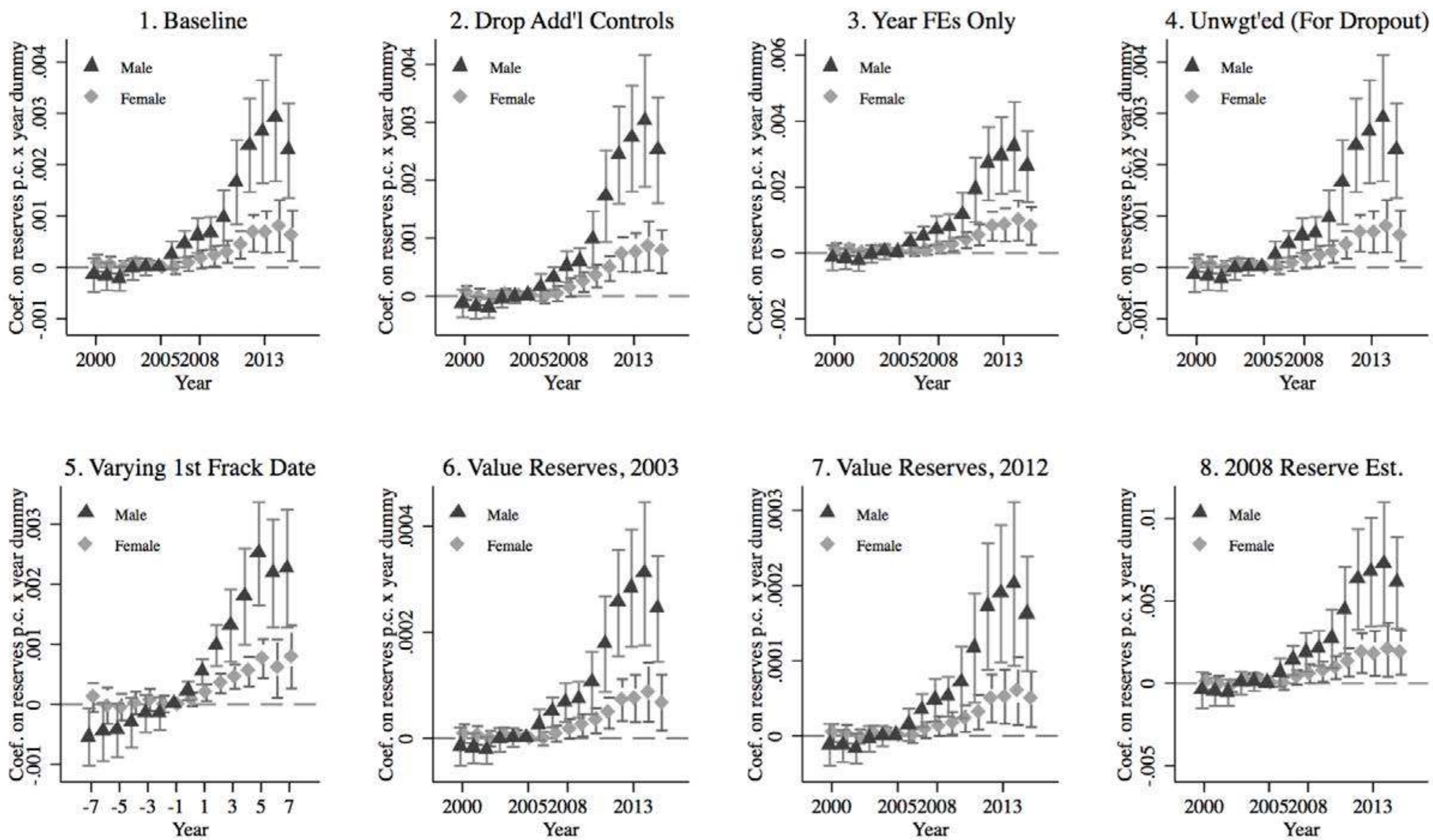
Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 5. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data in Panel A are from the Common Core of Data from 1999-00 to 2014-15, and enrollment ratio is defined as the ratio of 11th and 12th grade enrollment to the 17-18-year-old population. Sample is limited to 200 CZs in the 14 analysis states.

Appendix Figure A2 -
Table 5 Specification Checks, Dropout Rate



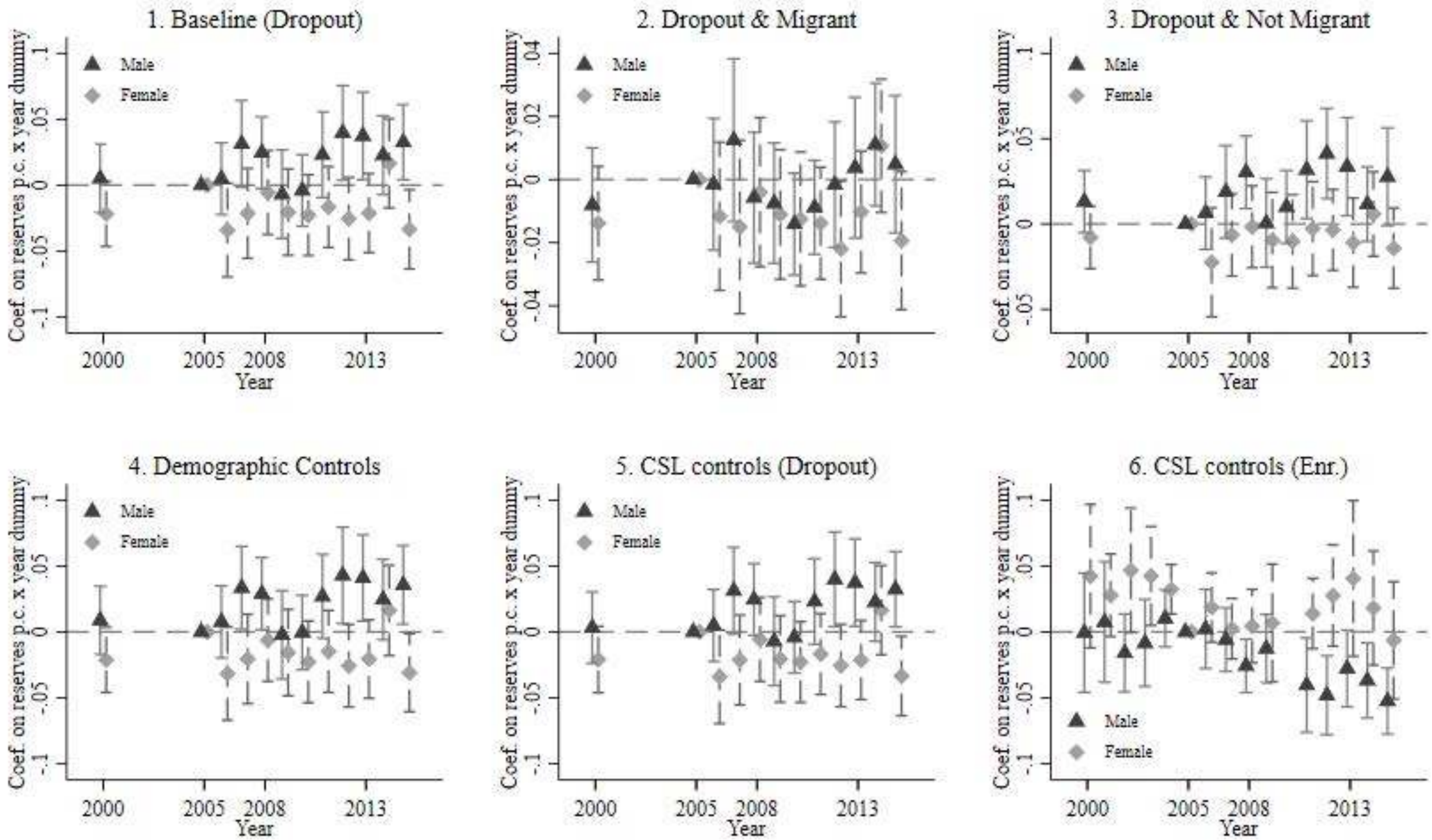
Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 5. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from the 2000 Census and the 2005-2015 ACS PUMS, and dropout is defined as being not currently enrolled and without a high school degree. Sample is limited to 202 CZs in the 14 analysis states.

Appendix Figure A3 - Table 5 Specification Checks, ln(Expected Monthly Earnings)



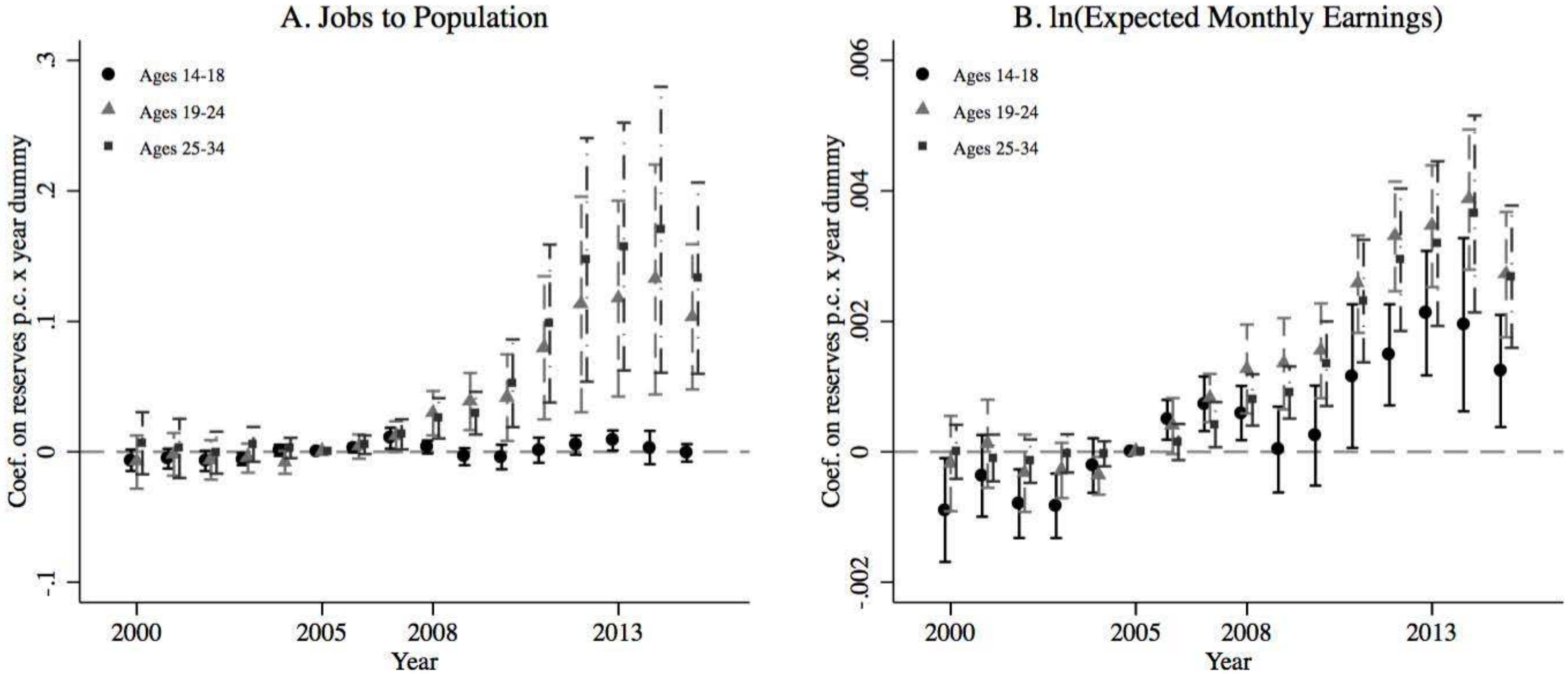
Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 5. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from the Quarterly Workforce Indicators (Panel B) and span 2000-2015; see Data Appendix. Sample is limited to 202 CZs in the 14 analysis states.

Appendix Figure A4 -
Table 6 Robustness Checks



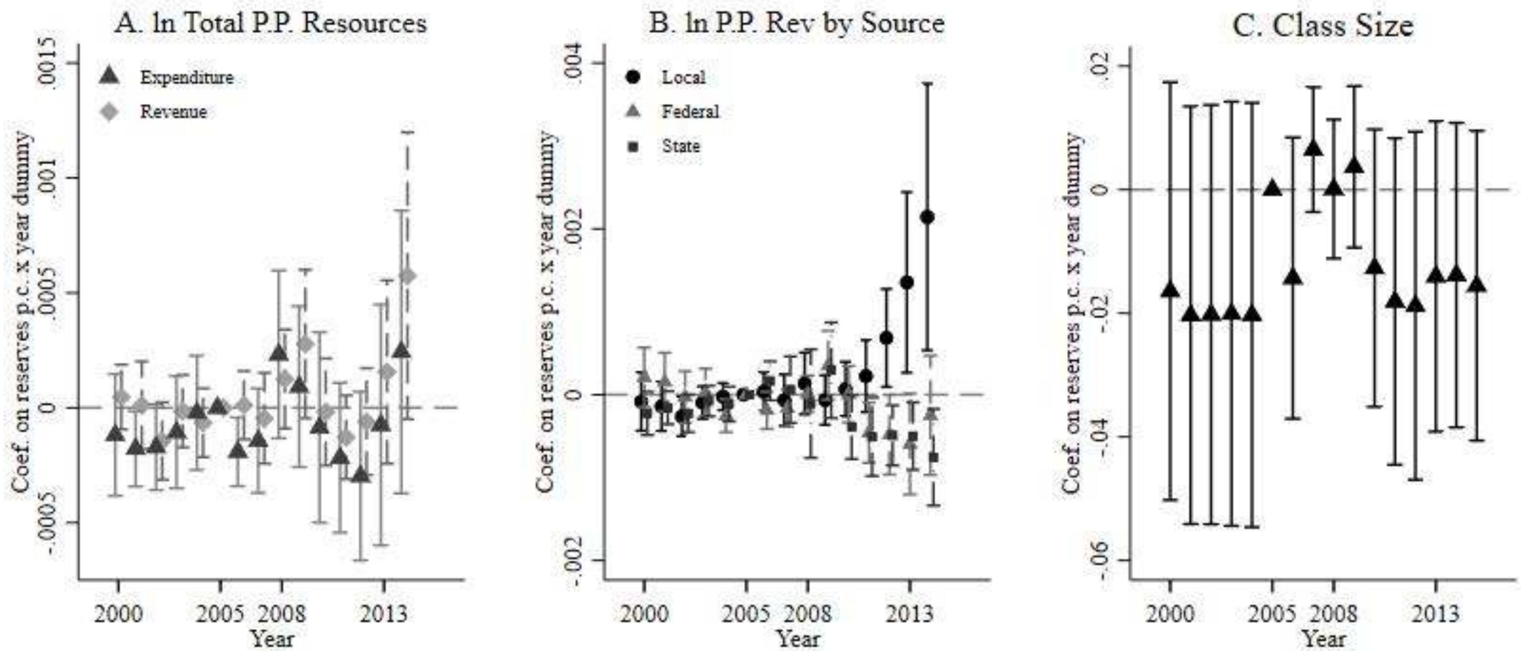
Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 6. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data in Panels 1-5 are from the 2000 Census and the 2005-2015 ACS PUMS, and dropout is defined as being not currently enrolled and without a high school degree. Data in Panel 6 are from the Common Core of Data from 1999-00 to 2014-15, and enrollment ratio is defined as the ratio of 11th and 12th grade enrollment to the 17-18-year-old population. See Data Appendix. Sample is limited to 202 CZs (Panels 1-5) or 200 CZ (Panel 6) in the 14 analysis states.

Appendix Figure A5 -
Table 7 Male Jobs and Expected Earnings by Age



Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 7. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from the Quarterly Workforce Indicators (Panel B) and span 2000-2015; see Data Appendix. Sample is limited to 202 CZs in the 14 analysis states.

Appendix Figure A6 -
Table 8 School Resources



Notes: Graphs plot coefficients on interactions between year dummies and predicted shale oil and gas reserves per capita (measured in 1000s of MMBTUs and normalizing by year 2000 CZ population) (omitting the interaction with the 2005 dummy for identification) from regressions that correspond with each of the specifications presented in Table 8. Inference is robust to heteroskedasticity and error correlation within commuting zones over time. Capped vertical lines represent 90 percent confidence intervals on the coefficient estimates. Data are from the Census of Governments and Annual Surveys of State and Local Government Finances (Panels A and B) and the Common Core of Data (Panel C) and span 2000-2015. Sample is limited to 202 CZs in the 14 analysis states.

Appendix Table A1 -
The Effect of Shale Oil and Gas Reserves and the Introduction of Fracking on Oil and Gas Production

	Per-capita Shale Oil and Gas Production				Per-capita Conventional Oil and Gas Production	
Dependent Variable:	in 1000s of MMBTUs	Inverse Hyperbolic Sine of MMBTUs	in 1000s of real $2012	Inverse Hyperbolic Sine of value	in 1000s of real $2012	in 1000s of MMBTUs
	(1)	(2)	(3)	(4)	(5)	(6)
Average effect 2011-15, *reserves*>0	1.58	0.26	18.10	0.41	0.01	-0.02
	(0.80)	(0.09)	(9.76)	(0.12)	(0.08)	(0.53)
Coefficient on:						
Shale reserves per capita	0.0060	0.0018	0.0764	0.0037	-0.0002	0.0059
x 2006-10	(0.0045)	(0.0008)	(0.0648)	(0.0011)	(0.0010)	(0.0062)
Shale reserves per capita	0.0375	0.0062	0.4303	0.0097	0.0002	-0.0004
x 2011-15	(0.0190)	(0.0021)	(0.2319)	(0.0029)	(0.0020)	(0.0125)
p : = coefs. (across yr. groups)	*0.039*	*0.007*	*0.043*	*0.008*	*0.759*	*0.626*
Observations	3,232	3,232	3,232	3,232	3,232	3,232
R-squared	0.616	0.799	0.661	0.880	0.891	0.935

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year. Data on oil and gas production by well type are from DrillingInfo from 2000-2015 and are converted into 2012 dollars using the energy CPI and MMBtus using 2012 conversion factors reported by EIA. All regressions include state-by-year fixed effects, commuting zone fixed effects, and time-varying effects of each of the 2000 Census commuting zone characteristics listed in Table 1 Panel D.

Appendix Table A2 -
The Effect of Shale Oil and Gas Reserves and the Introduction of Fracking on Jobs to Population, by Sex and Educational Attainment: Coefficients for Table 3

	Jobs (by Education) to Group Population, Ages 25+			
	High Sch.	High Sch.	College	College
Dependent variable:	Dropouts	Graduates	Attendees	Graduates
	(1)	(2)	(3)	(4)
		A. Men (N=3,232)		
Shale reserves per capita	0.0162	0.0185	0.0245	0.0067
x 2006-10	(0.0098)	(0.0086)	(0.0128)	(0.0029)
Shale reserves per capita	0.1666	0.1096	0.1440	0.0385
x 2011-15	(0.0697)	(0.0471)	(0.0589)	(0.0160)
p : = coefs. (across yr. groups)	*0.015*	*0.020*	*0.011*	*0.021*
R-squared	0.902	0.864	0.893	0.964
		B. Women (N=3,232)		
Shale reserves per capita	-0.0042	0.0021	0.0100	-0.0014
x 2006-10	(0.0099)	(0.0035)	(0.0052)	(0.0027)
Shale reserves per capita	0.0146	0.0131	0.0316	0.0036
x 2011-15	(0.0177)	(0.0068)	(0.0082)	(0.0042)
p : = coefs. (across yr. groups)	*0.055*	*0.017*	*0.000*	*0.042*
R-squared	0.973	0.953	0.971	0.975
		C. Male-Female Difference (N=6,464)		
Shale reserves per capita	0.0204	0.0164	0.0145	0.0081
x 2006-10	(0.0136)	(0.0082)	(0.0114)	(0.0021)
Shale reserves per capita	0.1520	0.0965	0.1123	0.0349
x 2011-15	(0.0664)	(0.0420)	(0.0533)	(0.0129)
p : = coefs. (across yr. groups)	*0.019*	*0.022*	*0.024*	*0.028*
p : = coefs. (across sex, 2011-15)	*0.023*	*0.023*	*0.036*	*0.008*
R-squared	0.933	0.887	0.927	0.969

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data on jobs for individuals ages 25 and over are from the 2000-2015 QWI and correspond to unweighted averages of beginning of quarter figures reported throughout the year; data on the population ages 25 and over are from SEER; and estimates of education shares in the population ages 25-64 are from the Census and ACS. Ratios multiply group-specific jobs by group-specific population, estimated as the product of population and the relevant group share in the population. Cell entries give coefficients (standard errors) from model 2 (Panels A and B) or a fully-interacted version of model 2 using pooled data (Panel C). All regressions include state-by-year and CZ fixed effects and time-varying effects of the CZ observables summarized in Table 1 Panel D. Each CZ is given equal weight in the estimation. Standard errors clustered on CZ are in parentheses.

Appendix Table A3 -
The Effect of Shale Oil and Gas Reserves and the Introduction of Fracking on Expected Monthly Earnings, by Sex and Educational Attainment: Coefficients for Table 3

	ln(Expected Monthly Earnings)			
	High Sch.	High Sch.	College	College
Dependent variable:	Dropouts	Graduates	Attendees	Graduates
	(1)	(2)	(3)	(4)
		A. Men (N=3,232)		
Shale reserves per capita	0.0009	0.0008	0.0007	0.0004
x 2006-10	(0.0003)	(0.0003)	(0.0003)	(0.0002)
Shale reserves per capita	0.0030	0.0027	0.0025	0.0016
x 2011-15	(0.0008)	(0.0006)	(0.0006)	(0.0006)
p : = coefs. (across yr. groups)	*0.000*	*0.000*	*0.000*	*0.004*
R-squared	0.942	0.932	0.939	0.962
		B. Women (N=3,232)		
Shale reserves per capita	0.0003	0.0002	0.0002	-0.0001
x 2006-10	(0.0002)	(0.0001)	(0.0001)	(0.0001)
Shale reserves per capita	0.0012	0.0007	0.0007	0.0002
x 2011-15	(0.0004)	(0.0003)	(0.0002)	(0.0002)
p : = coefs. (across yr. groups)	*0.001*	*0.004*	*0.000*	*0.026*
R-squared	0.971	0.942	0.954	0.968
		C. Male-Female Difference (N=6,464)		
Shale reserves per capita	0.0006	0.0006	0.0005	0.0005
x 2006-10	(0.0002)	(0.0002)	(0.0002)	(0.0001)
Shale reserves per capita	0.0018	0.0020	0.0018	0.0014
x 2011-15	(0.0006)	(0.0004)	(0.0004)	(0.0004)
p : = coefs. (across yr. groups)	*0.008*	*0.000*	*0.000*	*0.004*
p : = coefs. (across sex, 2011-15)	*0.002*	*0.000*	*0.000*	*0.001*
R-squared	0.973	0.964	0.955	0.970

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 95th percentile of size (based on 2000 population) for their state (202 of 259 CZs). The unit of observation is CZ-year-sex. Data on monthly earnings for individuals ages 25 and over are from the 2000-2015 QWI and correspond to unweighted averages of beginning of quarter figures reported throughout the year. Expected monthly earnings multiply reported monthly earnings by the group-specific jobs-to-population ratio. Cell entries give coefficients (standard errors) from model 2 (Panels A and B) or a fully-interacted version of model 2 using pooled data (Panel C). All regressions include state-by-year and CZ fixed effects and time-varying effects of the CZ observables summarized in Table 1 Panel D. Each CZ is given equal weight in the estimation. Standard errors clustered on CZ are in parentheses.

Appendix Table A4 -
Sensitivity of the Estimates for Dropout, Enrollment, and Earnings to Choice of Specification
Coefficients for Table 5, Panel B

Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
			No time-varying effects of CZ observables			
Shale reserves per capita	0.0002	0.0002	0.0040	0.0053	0.0006	0.0001
x 2006-10	(0.0136)	(0.0123)	(0.0087)	(0.0079)	(0.0002)	(0.0001)
Shale reserves per capita	-0.0297	0.0088	0.0269	0.0083	0.0026	0.0007
x 2011-15	(0.0150)	(0.0148)	(0.0103)	(0.0087)	(0.0006)	(0.0002)
p : = coefs (across yr. groups)	*0.129*	*0.449*	*0.003*	*0.678*	*0.000*	*0.000*
R-squared	0.746	0.759	0.496	0.450	0.932	0.949
			Year fixed effects			
Shale reserves per capita	-0.0049	-0.0031	0.0063	0.0020	0.0008	0.0001
x 2006-10	(0.0149)	(0.0135)	(0.0074)	(0.0073)	(0.0003)	(0.0001)
Shale reserves per capita	-0.0355	0.0003	0.0349	0.0135	0.0028	0.0008
x 2011-15	(0.0146)	(0.0189)	(0.0096)	(0.0076)	(0.0007)	(0.0003)
p : = coefs (across yr. groups)	*0.094*	*0.779*	*0.000*	*0.106*	*0.000*	*0.004*
R-squared	0.715	0.732	0.478	0.419	0.927	0.941
			Unweighted Estimates (Changes Dropout Only Relative to Baseline)			
Shale reserves per capita	-0.0035	-0.0024	0.0107	-0.0105	0.0007	0.0001
x 2006-10	(0.0163)	(0.0166)	(0.0038)	(0.0060)	(0.0002)	(0.0001)
Shale reserves per capita	-0.0340	0.0085	0.0244	-0.0013	0.0025	0.0006
x 2011-15	(0.0151)	(0.0192)	(0.0055)	(0.0053)	(0.0006)	(0.0002)
p : = coefs (across yr. groups)	*0.065*	*0.385*	*0.00*	*0.07*	*0.00*	*0.00*
R-squared	0.764	0.777	0.360	0.373	0.936	0.954

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for each outcome except the enrollment-to-population ratio, where there are 200 CZs). The unit of observation is CZ-year-sex. Unless otherwise noted (Panel B), all regressions include state-by-year and CZ fixed effects, as well as time-varying effectsd of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 as described in the text. Regressions for dropout rates (columns 3 and 4) are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean dropout rates. Standard errors (in parentheses) are clustered on CZ.

Appendix Table A5 -
Sensitivity of the Estimates for Dropout, Enrollment, and Earnings to Choice of Specification
Coefficients for Table 5, Panel C

Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
			Use first frack date for largest play in state			
Shale reserves per capita	-0.0239	0.0053	0.0162	0.0105	0.0018	0.0005
x post-fracking	(0.0108)	(0.0162)	(0.0072)	(0.0068)	(0.0005)	(0.0002)
R-squared	0.763	0.776	0.538	0.494	0.934	0.954
			Long difference: 2000 versus 2011-15			
Shale reserves per capita	-0.0334	0.0024	0.0248	0.0057	0.0025	0.0006
x 2011-15	(0.0265)	(0.0159)	(0.0097)	(0.0083)	(0.0007)	(0.0003)
R-squared	0.795	0.808	0.679	0.629	0.925	0.945

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for each outcome except the enrollment-to-population ratio, where there are 200 CZs). The unit of observation is CZ-year-sex. Unless otherwise noted (Panel B), all regressions include state-by-year and CZ fixed effects, as well as time-varying effectsd of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 as described in the text. Regressions for dropout rates (columns 3 and 4) are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean dropout rates. Standard errors (in parentheses) are clustered on CZ.

Appendix Table A6 -
Sensitivity of the Estimates for Dropout, Enrollment, and Earnings to Choice of Specification:
Coefficients for Table 5, Panel D

Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
	Value of reserves using 2003 prices (1000s of $2012)					
Shale reserves per capita	-0.0003	-0.0002	0.0006	-0.0002	0.0001	0.0000
x 2006-10	(0.0017)	(0.0018)	(0.0009)	(0.0008)	(0.0000)	(0.0000)
Shale reserves per capita	-0.0036	0.0009	0.0029	0.0004	0.0003	0.0001
x 2011-15	(0.0016)	(0.0021)	(0.0009)	(0.0008)	(0.0001)	(0.0000)
p : = coefs (across yr. groups)	*0.06*	*0.40*	*0.01*	*0.49*	*0.00*	*0.00*
R-squared	0.764	0.776	0.539	0.493	0.938	0.955
	Value of reserves using 2012 prices (1000s of $2012)					
Shale reserves per capita	0.0000	0.0003	0.0004	0.0001	0.0001	0.0000
x 2006-10	(0.0011)	(0.0012)	(0.0004)	(0.0005)	(0.0000)	(0.0000)
Shale reserves per capita	-0.0023	0.0008	0.0019	0.0006	0.0002	0.0000
x 2011-15	(0.0011)	(0.0014)	(0.0006)	(0.0005)	(0.0001)	(0.0000)
p : = coefs (across yr. groups)	*0.05*	*0.56*	*0.01*	*0.37*	*0.00*	*0.01*
R-squared	0.764	0.776	0.538	0.493	0.937	0.955
	Simulated reserves in 2008 (1000s of MMBTUs)					
Shale reserves per capita	0.0395	0.0111	0.0034	-0.0148	0.0020	0.0005
x 2006-10	(0.0353)	(0.0506)	(0.0251)	(0.0181)	(0.0008)	(0.0004)
Shale reserves per capita	-0.0605	0.0409	0.0536	0.0155	0.0064	0.0017
x 2011-15	(0.0543)	(0.0570)	(0.0251)	(0.0202)	(0.0020)	(0.0008)
p : = coefs (across yr. groups)	*0.07*	*0.53*	*0.04*	*0.11*	*0.00*	*0.02*
R-squared	0.763	0.776	0.538	0.493	0.933	0.954

Notes: Underlying sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for each outcome except the enrollment-to-population ratio, where there are 200 CZs). The unit of observation is CZ-year-sex. Unless otherwise noted (Panel B), all regressions include state-by-year and CZ fixed effects, as well as time-varying effectsd of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Cell entries give estimated effects on outcomes as of 2011-15 for the average CZ in the estimation sample with any shale reserves, calculated from model 2 as described in the text. Regressions for dropout rates (columns 3 and 4) are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean dropout rates. Standard errors (in parentheses) are clustered on CZ.

Appendix Table A7 -
Additional Specification Checks on Sample Selection Criteria

Dependent variable:	Enrollment ratio, Gr 11-12: Ages 17-18		Dropout, Ages 17-18		ln(Expected Earnings) Ages 25+	
	Male	Female	Male	Female	Male	Female
	(1)	(2)	(3)	(4)	(5)	(6)
Using all 17 States with Major Shale Plays (Does Not Change Earnings Estimates Relative to Baseline)						
Average effect 2011-15, *reserves*>0	-1.25	0.34	0.98	0.04	0.104	0.026
	(0.58)	(0.78)	(0.31)	(0.33)	(0.027)	(0.010)
Shale reserves per capita	-0.0017	-0.0014	0.0038	-0.0053	0.0007	0.0001
x 2006-10	(0.0151)	(0.0159)	(0.0080)	(0.0075)	(0.0002)	(0.0001)
Shale reserves per capita	-0.0320	0.0087	0.0252	0.0010	0.0025	0.0006
x 2011-15	(0.0150)	(0.0200)	(0.0080)	(0.0084)	(0.0006)	(0.0002)
p : = coefs (across yr. groups)	*0.06*	*0.43*	*0.00*	*0.44*	*0.00*	*0.00*
p : = coefs. (across sex, 2011-15)	*0.13*		*0.03*		*0.000*	
R-squared	0.761	0.778	0.524	0.479	0.936	0.954
Dropping Smallest & Largest 10% of CZs in State						
Average effect 2011-15, *reserves*>0	-1.65	0.30	0.98	0.22	0.108	0.026
	(0.56)	(0.82)	(0.40)	(0.31)	(0.026)	(0.010)
Shale reserves per capita	-0.0044	-0.0071	0.0012	0.0019	0.0007	0.0001
x 2006-10	(0.0158)	(0.0162)	(0.0088)	(0.0072)	(0.0002)	(0.0001)
Shale reserves per capita	-0.0388	0.0070	0.0229	0.0052	0.0025	0.0006
x 2011-15	(0.0131)	(0.0192)	(0.0094)	(0.0072)	(0.0006)	(0.0002)
p : = coefs (across yr. groups)	*0.05*	*0.20*	*0.00*	*0.66*	*0.000*	*0.001*
p : = coefs. (across sex, 2011-15)	*0.07*		*0.12*		*0.000*	
R-squared	0.794	0.802	0.545	0.503	0.942	0.957
Dropping Smallest & Largest 5% of CZs in State						
Average effect 2011-15, *reserves*>0	-1.39	0.30	0.79	0.47	0.105	0.026
	(0.59)	(0.78)	(0.38)	(0.32)	(0.026)	(0.010)
Shale reserves per capita	-0.0027	-0.0009	-0.0020	-0.0028	0.0007	0.0001
x 2006-10	(0.0158)	(0.0157)	(0.0082)	(0.0066)	(0.0002)	(0.0001)
Shale reserves per capita	-0.0342	0.0073	0.0194	0.0116	0.0026	0.0006
x 2011-15	(0.0145)	(0.0191)	(0.0093)	(0.0079)	(0.0006)	(0.0002)
p : = coefs (across yr. groups)	*0.05*	*0.52*	*0.00*	*0.09*	*0.000*	*0.001*
p : = coefs. (across sex, 2011-15)	*0.11*		*0.49*		*0.000*	
R-squared	0.770	0.783	0.561	0.510	0.941	0.959

Notes: Baseline sample consists of the 202 CZs in the 14 of the lower 48 states with any major shale gas or oil play (17 states), with available data from the QWI for 2000-2015, and between the 5th and 90th percentile of size (based on average population of 17-18-year-olds over 2000-2005) for their state (202 of 259 CZs for each outcome except the enrollment-to-population ratio, where there are 200 CZs). Panels change this estimation sample in various ways, adding3 states with major shale plays but without QWI data for 2000 (239 CZs), trimming the top and bottom 10% of CZs within each state in the original 14-state sample (186 CZs), and trimming the top and bottom 5% of CZs within each state in the original 14-state sample (214 CZs). The unit of observation is CZ-year-sex. All regressions include state-by-year and CZ fixed effects, as well as time-varying effectsd of each of the 2000 Census CZ characteristics listed in Table 1 Panel D. Regressions for dropout rates (columns 3 and 4) are weighted by the number of Census or ACS respondents used to generate the CZ-year-sex mean