The Impact of the WWI Agricultural Boom and Bust on Female Opportunity Cost and Fertility

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Abstract:

Using variation in crop prices induced by large swings in demand surrounding World War I, we examine the fertility response to increases in crop revenues during the period 1910-1930. Our estimates from samples utilizing both complete count decennial census microdata and newly collected county-level data from state health reports indicate that a doubling of the agricultural price index reduced fertility by around 10-13 percent throughout the period. This effect persists years after the collapse of the war boom. We further show that fertility declines along both the extensive and intensive margins, which is consistent with predictions from models of rising opportunity costs. Our findings are robust to many potential confounds.

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1. Introduction

A recurring theme in economics and demography is that as economies develop, a demographic transition will occur, whereby declining fertility rates accompany rising incomes (Galor, 2005; Jones and Tertilt 2009; Galor and Weil, 2000, Brueckner and Schwandt, 2015). Theory makes sharp predictions regarding the relationship between fertility, income shocks, and changes in the opportunity cost of time (Becker, 1960; Becker and Lewis; 1973; Becker and Tomes, 1976; Rosenzweig, 1977), however, empirical evidence is needed to understand whether pure income shocks or changes in the opportunity costs of time dominate in specific contexts. A recent literature has developed that takes advantage of exogenous variation in income, employment, or wealth to identify their impact on fertility.[4] The recent literature has broadly shown that increases in the male wage or income leads to increases in fertility, while increasing female wages result in reductions in fertility. Most of this evidence comes from the period including or following the Baby Boom. Scholars know relatively less about the causal factors contributing to the decline from the late 19th Century through about 1930.

In this paper, we draw on rich, historical data from the United States to understand how fertility responds to aggregate shocks that change both incomes and the relative opportunity costs of women. Specifically, we take advantage of a period of significant agricultural price variation, the agricultural commodity boom and bust in the United States surrounding World War I (WWI). During the period we study, 1910-1930, both the General Fertility Rate and Crude Birth rate fell by approximately 29 percent, which is as large a decline as the Baby Boom was a boom. The agricultural price variation induced by international events combined with large changes in fertility make it an ideal setting to explore the link between income shocks and fertility. The agricultural commodity boom in WWI was entirely unexpected as fields in Europe were destroyed. Additionally, the magnitude of the shock was massive. U.S. agricultural exports doubled in the second half of the 1910s, in some cases prices more than doubled, and agricultural receipts increased by 70 percent (Henderson, Gloy, and Boehlje, 2011). Farmers

[4] See for example Lindo, 2010; Ananat, Gassman-Pines, and Gibson-Davis, 2013; Huttunen and Kellokumpu, 2017; Black et al., 2013; Yonzan, Timilsina, and Kelly, 2020; Fishback, Haines, and Kantor, 2007; Lovenheim and Mumford, 2013; Dettling and Kearney, 2014; Bailey and Collins, 2011; Lewis, 2018; Fujii and Shonchoy, 2020; Wanamaker, 2012; Schaller, 2016.

expected the boom to persist, as documented by rising land prices.[5] The subsequent bust following the Treaty of Versailles also came as a surprise as Europe rapidly recovered post war.[6] We document a close link between female agricultural wages and the crop price variation, which we argue makes changes in female opportunity costs the driving mechanism for the fertility response.

The first few decades of the 20th Century United States are understudied due to limited data availability. The federal government did not begin recording births until 1915 with the creation of the Birth Registration Area (BRA), which was not complete until Texas joined in 1933. To gain insight in the pre-war period, we digitize annual county-level birth tabulations from available state health reports prior to a state's entry into the BRA to push back the series to 1910 for thirty-two states. The longer panel enables us to capture any variation in pre-trends that might confound our estimate. Additionally, we turn to the complete count Population Census for the years 1910, 1920, and 1930 to compare birth outcomes for women differentially exposed to the agricultural boom and bust, netting out common locational and cohort effects.

To identify the empirical relationship between changes in fertility and changes in agricultural income, we complement our birth data with a measure of annual county-level agricultural crop revenue. Specifically, we augment the approach of Rajan and Ramcharan (2015) and Jaremski and Wheelock (*forthcoming*) to construct a county-level agricultural price index. The index combines pre-war crop production bundles at the county-level with aggregate crop specific price shocks to generate our key source of spatial-temporal variation. We use the index as our measure of agricultural income because annual county-level crop receipts and gender specific wages are otherwise unavailable consistently at a disaggregated level. We then estimate the relationship between the agricultural price index and fertility, controlling for a rich set of covariates, location specific fixed effects, and time fixed effects. Under the assumption that that the

[5] Agricultural land prices increased an average of 170 percent nationally by 1920 relative to prewar prices (Table 540, USDA Yearbook of Agriculture, 1931). In places such as Arkanasas, Georgia, Mississippi, North Carolina, and South Carolina, average land prices increased between 217-230 percent per acre relative to their prewar levels.

[6] In the years following WWI, significant price volatility remained, at least in part as a result of increased international competition and new domestic policy (i.e., Capper-Volstead Act, 1922, Fordney-McCumber Tariff, 1922).

initial crop specialization patterns are driven by agro-climatic variables and that national crop specific prices are driven by international events and climate shocks, variation in the agricultural price index is exogenous to the measures of birth and fertility, thus permitting a causal interpretation for our estimates. We also implement an instrumental variables strategy exploiting crop suitability measures to generate predicted crop shares and find a similar pattern of results.

Our estimates consistently show that the agricultural boom reduced fertility in the short-run, measured by annual county-level births, as well as in the medium and long-run, measured by counts of children from the population census. In the short-run, we estimate that annual county-level births fell by 13-16 percent when crop prices doubled. Evaluated at the average value of the agricultural index, agricultural crop variation led to a 3.7 percent decline in births which explains about 12.75 percent of the overall decline in fertility between 1910 and 1930. This estimate is robust to a variety of potential confounds, including county-level WWI induction rates, exposure to the Spanish Influenza Pandemic, and Prohibition controls. Turning to the complete count census samples, we estimate that the number of children under the age of 5 for women in prime child bearing years fell 9 percent evaluated at the mean. Using a proxy for long-run fertility, we estimate a 3.7-4.8 percent relative decline in the total number of children in the home, indicating that fertility was not merely delayed or retimed. We then test for changes in both the extensive and intensive margins of fertility. Here we estimate that changes at the extensive margin explain one-third of the net decline while changes at the intensive margin explain two-thirds. These findings are consistent with theoretic predictions found in Galor (2012) and Aaronson, Lange, and Mazumder (2014), who highlight the effects of increasing the price of child quantity on fertility.

In the full draft, we consider several mechanisms including delayed marriage, the differential effects by ownership, capital adoption, the labor intensity of agriculture, and implement an instrumental variables strategy based on agricultural suitability.

Our paper contributes to multiple strands of literature. First, our paper complements the body of work that seeks to understand the causal relationship between economic shocks and fertility. Our paper fills a gap in the literature regarding demographic changes within the United States prior to the Baby Boom. Relative to recent work by Ager, Herz, and Brueckner (*forthcoming),* who examine the impact of the Boll Weevil on structural transformation and fertility in the American South, our paper examines the broader impacts of agricultural price volatility on family structure and brings several new data sources to light. Second, we view our work as contributing to the broader modern development economics literature that relates agricultural and natural resource shocks on families and family structure (Beegle, Dehejia, and Gatti, 2006; Kruger, 2007; Akresh, 2009; Cogneau and Jedwab, 2012). Our estimates show that agricultural commodity price shocks can also affect the size of the family, which closely relates to Schultz (1985) and Corno, Hildebrandt, and Voena (*forthcoming*). Finally, our work contributes to a recent historical literature that seeks to understand the economic impacts of WWI, joining a growing body of work that explores family formation in Europe (Abramitzky, Delavande, and Vasconcelos, 2011; Vandenbrouke, 2014; Gay, 2019; and Boehnke and Gay, 2020).

2. World War I Agricultural Boom and Bust

The 1910s and 1920s were a period of significant volatility within American agriculture. WWI created an unprecedented demand shock for American agricultural products. As war raged in Europe, wheat production fell by over 50 percent in both France and Italy, oat production dropped 59 percent in Germany, and livestock plummeted to a quarter of their prewar levels in Denmark (Nourse, 1924). Distress in ocean shipping further increased demands for US products that were less exposed on ocean shipping lanes than goods originating from more distant sources, such as Argentina and Australia. In response, prices for American goods rose sharply, with significant crop specific variation.

Because of the rapid increase in agricultural prices, production on the home front ramped up. Between 1914 and 1919, 30 million new acres of land were brought into production, representing a 9 percent relative increase (Olmstead and Rhode, 2006).Given the sharp increase in prices and the increased production, the aggregate value of crops harvested in the United States more than doubled over the same time horizon (Acquaye, Alston, and Pardey, 2006). The rapid increase in production strained input markets. Labor was particularly scarce in rural areas as people flocked to cities to take advantage of the wartime manufacturing boom. While America's war effort was modest in WWI relative to World War II in terms of manpower, over 4 million young men were drafted to become doughboys, further tightening the labor market. Ultimately, these factors led to higher wages in agriculture.

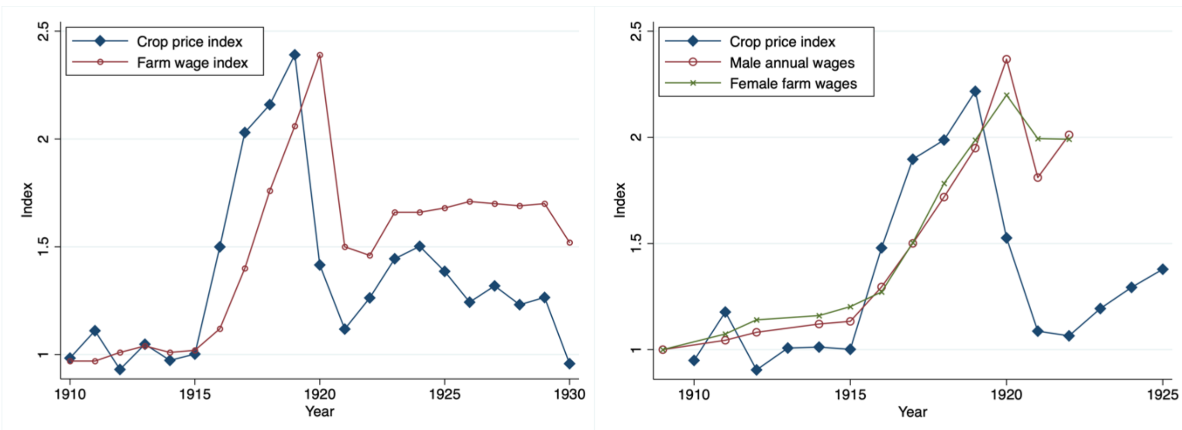
How would higher agricultural wages affect fertility? Following models presented in Galor (2012) and Aaronson, Mazumder, and Lange (2014), individuals that maximize utility over consumption and children while investing in child quality will adjust their optimal fertility if either the fixed cost of child rearing changes or if the cost of child quality changes. Factors that affect the fixed cost of raising children independent of quality investments unambiguously lead to declines in fertility along both the extensive and intensive margins. In our context, increasing labor tightness that leads to higher market wages would increase the fixed cost of raising children and should result in declines in fertility along both margins of adjustment.

Ideally, we would observe gender specific county-level wages annually and relate changes in wages or income to fertility along both the extensive and intensive margins. However, no data exist with national coverage that enable us to test the direct link between gender specific wages and changes in fertility. While we do not generally have disaggregated wage data, we have two key sources that suggest a close link between wages and agricultural receipts. In Figure 1, Panel A, we report the average annual crop price index that we construct and describe Section 3. We also plot the national average farm labor wage index from 1910-1930 (USDA, 1931). The USDA labor index closely tracks the crop index prior to WWI. In 1919, at the peak of the crop index, the wage index doubled from its prewar levels, remaining elevated through the price index bust in 1920. Throughout the remainder of the 1920s, agricultural wages remained 50 to 70 percent above their prewar levels.

Our second piece of evidence comes from Pennsylvania. The Pennsylvania Department of Agriculture reported the weekly wage of women hired as domestic help on farms with board and the annual wage with board for males at the county-level between 1908 and 1922. In Figure 1 Panel B, we construct the crop price index for

Pennsylvania. In Pennsylvania, the crop index reached a peak of 2.3 in 1919, while the male and female wage indices peak in 1920 and remain elevated through 1922 when the series ends. Female wages increased by a factor of 2.2, while male wages increased by a factor of 2.37. As with the national wage data, wages in Pennsylvania remain elevated in the post war years, 1920-1922. These data highlight a strong link between our measured crop index and female wages.

Figure 1 - Agricultural Price Index and Wage Index 1910-1930



Panel (A) Panel (B)

Notes: Panel A reports the average national crop index over our analysis timeframe. The index is a function of the county's baseline crop mix from 1908 – 1914 and national price fluctuations (Carter et al., 2006). The farm wage index is drawn from the USDA Yearbook of Agriculture (1931). In Panel B, we construct the average crop index in Pennsylvania. Additionally, we report indexed values for the weekly wage of women working on farms with board and the male annual wage, as reported by the Pennsylvania Department Agriculture's Crop and Livestock report 1906-1922.

Outside of Pennsylvania, there is anecdotal evidence of efforts undertaken to alleviate tight labor markets that also led to wage gains. Here we provide one example, the Women's Land Army. At its peak, the Women's Land Army recruited upwards of 20,000 women from colleges and universities work on farms while living in camps (SSA, 1942). In California, women working under the direction of the Women's Land Army earned a minimum wage of $2/day or the market wage, whichever was greater. Based on reports by the USDA, published in newspapers across the country in 1918, the wage paid to the Women's Land Army was equivalent to the male daily wage that included room and board. To the extent that there were real wage increases for women in agriculture, this would tend to increase the opportunity cost of child bearing, at least in the short-run.

The agricultural boom in the US was short-lived and was followed by an abrupt and unexpected bust. The fields of Europe recovered quickly following the signing of the Treaty of Versailles. For example, Buyst and Franaszek (2010) report that crop specific yields recovered for most of Europe by 1922. Even Russia, in the midst of a civil war, was able to increase its agricultural output to pre-war levels by the mid-1920s (Markevich and Harrison, 2011). Following the end of WWI, agricultural commodity prices and wages remained elevated while the vast expansion in agriculture would be the source of foreclosures and financial hardship through the 1920s and 1930s (Alston, 1983; Rajan and Ramcharan, 2015).

3. Data

To estimate the relationship between the agricultural price volatility and fertility, we combine data from the county-level agricultural Census with two different samples detailing fertility. First, we construct a newly digitized dataset of county-level birth counts that predate the Federal Birth Registration Area (BRA), sourced from state health reports between 1910 and 1930 for 32 states. Second, we turn to the complete count public use data from the US Population Census for the years 1910, 1920, and 1930 (Ruggles et al., 2020).

3.1 Population and Agricultural Data

To create the agricultural price index at the county-level, we augment the index suggested by Rajan and Ramcharan (2015) and Jaremski and Wheelock (*forthcoming*). We begin by collecting county-level output for the 12 crops (corn, wheat, oats, barley, rye, buckwheat, flaxseed, cotton, tobacco, Irish potatoes, sweet potatoes, and forage crops) from the 1910 Census of Agriculture (Haines, Fishback, and Rhode, 2018).[7] We then multiply each county's 1910 crop output $Q_{i,c,1910}$, by the crop's annual national price, $P_{i,t}$, drawn from Carter et al. (2006) to compute the annual county-level crop

[7] Rajan and Ramcharan (2015) and Jaremski and Wheelock (*forthcoming*) omit forage crops from their index. We include them here because hay and other field grasses account for over 50 percent of cattle feed inputs and are the largest share of acreage in the arid West (https://www.usdairy.com/news-articles/do-dairy-cows-eat-food-people-could-eat).

revenue. Finally, we normalize the annual county-level crop revenue by the average county-level revenue for the period 1908 and 1914, using the average crop price, $\overline{P_i}$.

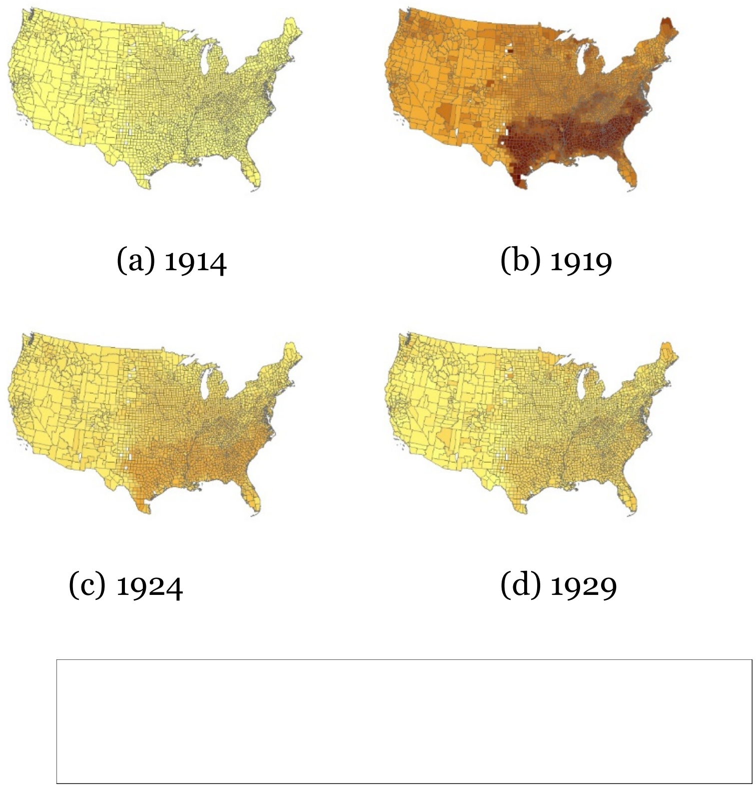
$$CropIndex_{c,t} = \frac{\sum_{i=1}^{12} Q_{i,c,1910} \times P_{i,t}}{\sum_{i=1}^{12} Q_{i,c,1910} \times \overline{P_i}}$$

By fixing the output at the 1910 value, and using national prices, we ensure that the variation we exploit is exogenous to the decisions of local farmers. For instance, we do not have to be concerned with the potential of endogenous crop mixes in response to the movement in prices.

In Figure 1, we highlighted the time variation in the index. Beginning in 1915, the time when agriculture began to collapse in Europe, crop prices begin to rise dramatically, reaching a peak in 1918, with prices increasing by over 250 percent. Following the signing of the Treaty of Versailles in 1919, agricultural prices fell dramatically, yet remained above their pre-WWI level. While WWI was a significant source of crop price variation, there were also meaningful swings in crop prices throughout the 1920s. Given prewar agricultural production and crop specialization patterns, there was significant heterogeneity across space in terms of the local intensity of the agricultural boom. Cotton, Irish Potatoes, Tobacco, and Flaxseed all experienced price increases exceeding 300%. Thus, areas such as the Southeast, where cotton and tobacco are grown, experienced relatively larger shocks than the Midwest or West. Similarly, portions of Minnesota and North Dakota experienced relatively larger price spikes due to their production of flax. In Figure 2, Panels A-D, we highlight the spatial temporal variation in the normalized index for the years 1914, 1919, 1924, and 1929.

In addition to the data from the Census of Agriculture, we also merge several county-level economic and demographic variables from the 1910 population census (Haines, 2010). The primary variables we use include the population, percent non-white, percent urban, percent aged 6-14, percent illiterate, and the value of manufactured goods per capita.

Figure 2 - Spatial Variation in Agricultural Index



(a) 1914 (b) 1919

(c) 1924 (d) 1929

Notes: Four snapshots of the geographic variation in the crop index over our analysis timeframe. WWI caused demand shocks to be concentrated in different counties based on their preexisting crop composition. Data from the Agricultural Census and Carter et al. (2006).

3.2 Annual County Level Birth Data

Our first measure of fertility is comprised of county-level birth counts constructed from a combination of Birth Registration Area data and State Board of Health Reports sources. We begin with county-level birth data reported by the Federal Birth Registration Area (Eriksson, Niemesh, and Thomasson, 2018). The Federal Birth Registration Area (BRA) was formed in 1915 and included 10 states at its inception, primarily in the Northeast and upper Midwest. States joined the BRA following an application and certification process, whereby the US Census Bureau verified that the state in question accurately recorded 95 percent of births. The BRA was not complete until 1933, when Texas joined. To supplement these data, we collected and entered county-level birth data using a combination of state health department annual reports, state vital statistics annual reports, and state board of health monthly and quarterly bulletins published prior to a state's entry into the BRA back to 1910.

To appear in the sample, a state must not have missing data for more than 2 years between 1910 and its entry into the BRA. We are able to construct an annual county-level birth panel covering 32 states between 1910 and 1930 (see Appendix Figure 3).[8] In Appendix Table 2 we report the year that each state in our sample entered the BRA and the years for which we collected and coded the state-level reports.

The advantage of the county-level data are that they allow us to test for instantaneous responses in fertility associated with the crop index variation at high frequency. However, these data also have limitations. First, the states that appear in our sample tend to be located in the northeast, upper Midwest, and West. Thus, our county-level sample does not use the most extreme variation stemming from large swings in cotton and tobacco prices. Secondly, given the reporting accuracy requirements to enter the BRA, an obvious concern with the use of pre-BRA data is its reliability (i.e., measurement error). Finally, the birth counts do not allow us to examine the whether the net change in fertility is driven by intensive or extensive margin responses.

While there is a strong correlation between the county-level data and the Census data, there still may be a lingering concern that the reporting quality is suspect. Our estimates would be biased if measurement error is systematically correlated with movements in agricultural commodity prices. To alleviate these types of measurement concerns, our regressions include a variety of county-level covariates, county fixed effect, and year fixed effects.

Our county-level dataset consists of 33,374 county-year observations. In several specifications, we restrict this sample, dropping counties with populations above the 90th percentile of the 1910 population. We do this to ensure that our estimates are driven by changes in agricultural commodity prices and are not driven by other economic consequences of WWI such as industrial growth in urban centers.

[8] Each state set their own policy in regards to the tracking of vital statistics. In many cases, states did not pass enabling legislation early enough to begin the certification process to enter the Census Birth Registration Area or Death Registration Area during our sample period.

3.3 Individual Level Dataset

Given the limitations of the county-level dataset, we turn to the complete count microdata from the 1910, 1920, and 1930 Census (Ruggles et al., 2020). In this sample, we focus on women aged 14-49 who are spouses or householders and not living in group quarters. We exclude Hawaii and Alaska from the analysis. Across the three waves of the census, this amounts to over 49 million woman-year observations. In several specifications we will limit our attention to women who were more likely of child bearing age (under 35), of which there are approximately 25 million observations in the sample. As would be expected, women in the under 35 sample have more children under the age of 5 and less total children than the full sample. We will further consider similar population restrictions, dropping women living in the most populated counties based on 1910 population counts. The rural, under age 35 sample contains just over 12 million women. Importantly, restricting the sample to rural areas increases the probability that a woman is living on the farm from 25 percent in the under 35 sample to 46 percent in the rural sample. Women under the age of 35 living in less populated counties have more children and are less likely to be white, likely reflecting both the importance of child labor on farms and our pre-Great Migration period of analysis. Women aged 30-40 have more total children at home but fewer who are under the age of 5. The complete count data have the advantage of a very large sample size; however, we lose time variation given the infrequent observation.

4. Empirical Strategy

In what follows, we seek to estimate the reduced form relationship between a measure of fertility and the agricultural index. A naïve regression of births on income would likely produce biased results for two reasons. First, births may have a causal impact on income. This is particularly a concern in agriculture where children provide labor for the family. Second, omitted variables which affect both income and fertility could bias the estimates in a naïve regression; health and productivity are two examples of important characteristics which cannot be directly measured yet surely play a role in determining births and income.

To address these concerns, we exploit exogenous variation in the prices of different crops. The primary specification when using the county-level dataset derived from state health reports is:

$$LogBirths_{c,t} = \alpha + \phi Index_{c,t-1} + X'_{c,t}\beta + \tau_t + \sigma_c + \epsilon_{c,t}. \qquad (1)$$

The dependent variable is the log number of births in county c and year t. The primary coefficient of interest is ϕ which should capture the lagged effect of the previous year's crop index on county births. We use the lagged value of the index rather than the contemporaneous value given the nine-month gestation period. Year (τ) and county (σ) fixed effects account for national differences over time and stationary differences between counties, respectively. The vector X is comprised of interactions between the year fixed effects and various baseline county-level characteristics. These controls are measured in 1910 and include the fraction of the population who is non-white, between the ages of 6 and 14, illiterate, and living in urban areas with over 2500 residents. We also interact year fixed effects with county-level manufacturing output per capita in 1910. By interacting these baseline observable characteristics with year fixed effects, this specification should net out differential trends in these dimensions that may otherwise be picked up by the lagged crop index.

Estimating the impact of the crop index on births using decennial census data requires alternative specifications due to the timing of the data collection and variable availability. Ideally, we would be able to use a measure of completed fertility, however this variable is not available in either the 1920 or 1930 Census.[9] The primary dependent variables we use from the complete count data are the number of children under age 5, the total number of children in the household, and an indicator of any children in the household. Given the available data, our specification compares women of the same age across Census years, netting out time specific and county specific effects. Because the available variables represent the cumulative fertility response over multiple years, we assign each woman a cumulative or average crop index. When investigating the impact on children under the age of 5, we use the average county index over the last 5 years.

[9] We are primarily interested in how agricultural prices affected fertility, making the "children ever born" variable another relevant outcome. Unfortunately, this variable is only available in the 1910 and 1940 censuses.

When focusing on the total number of children or presence of children we use the average county index over the previous decade. The general specification when using the census data is:

$$Kids_{i,c,t} = \alpha + \phi AvgIndex_{i,c,t} + X'_{i,c,t}\beta + \tau_t + \sigma_c + \epsilon_{i,c,t}. \qquad (2)$$

The observations are at the mother (i), county of residence (c), and census year (t) level. The index variable is one of those described above and is assigned to each mother based on her age and county of residence. Census year fixed effects (τ) and county fixed effects (σ) are included, as are fixed effects for age, marital status, race, size of residence location (based on total population), and whether the mother lives on a farm. The dependent variable is either the number of children under age 5 at the time of the census, the woman's total number of kids at the time of the census, an indicator of any children in the household, or the number of children in the household conditional on having at least one child present. The latter two measures allow us to examine fertility responses along the extensive and intensive margins directly.

Both equation (1) and (2) are intended to address the same underlying question: how did agricultural commodity price variation affect fertility decisions? A priori, it is unclear whether the estimates from both specifications should point in the same direction. For instance, fertility could immediately fall during the wartime boom due to increased labor scarcity and increased opportunity cost of childbearing. However, at the war's conclusion, as labor demand subsides, and the supply constraint is lessoned as men return from Europe, women may simply retime children, resulting in no net change in total children in the long-run. Similarly, the sudden wealth shock could lead to an immediate increase in child bearing, while in the long-run there is no change in the total number of children born.

An underlying assumption in all of the regression models is that the baseline county-level crop share weights, $Q_{i,c,1910}$, used in the construction of the index, are exogenous. Recently, economists have raised concerns that the weights in shift-share variables may generate a source of bias (Goldsmith-Pinkham, Sorkin, and Swift, 2020). In our context, agro-climatic variables such as precipitation, temperature, soil type, and the biological pest environment likely drive agricultural productivity and crop choice decisions. In the full draft we also implement an IV strategy following Fiszbein (forthcoming).

5. Estimates

Below we present our estimates from several different empirical specifications. We proceed by discussing our estimates in the short-run by examining the county-year sample and the complete count sample that focuses on children under 5 in the household. We then discuss several robustness checks that address many of the potential confounds. After describing the short-run estimates, we outline our estimates over a longer time horizon (10 years), where it is possible for families to adjust along multiple dimensions. After highlighting our main empirical estimates, we explore some of the underlying mechanisms and heterogeneity along key dimensions.

5.1 Short-Run Fertility Effects of Agricultural Boom

In Table 1, we report our estimates from equation (1) and equation (2). In Panel A, we report the baseline results using the county-level birth sample. Panel B includes the estimates using the number of children under age 5 from the complete count data as the outcome of interest. In each case, we report standard errors in parentheses, clustered at the state level. Our estimates are robust to several different assumptions regarding standard errors.[10] Using the county-level sample, we find that doubling the agricultural price index results in a statistically significant 12.8 log point reduction in births. After dropping urban counties, the estimated decline in births increases to 14.4 log points. When we include our set of county level covariates in Columns 3 and 4, the magnitude our estimate remains similar, declining by 12.6 – 14.8 log points. Given an average of 435 births per rural county, our estimates suggest that there were between 55 and 64 fewer births per county-year given a doubling in agricultural prices. Evaluating at the average agricultural index value throughout the period, our estimates suggest a 3.7 percent reduction in fertility. Given the 29 percentage point decline in aggregate fertility, 1910-1930, the agricultural price volatility explains about 12.75 percent of the

[10] Clustering at the state level generates the most conservative estimates. Clustering at the county level or computing Conley (1999) standard errors for the county sample using a distance cutoff of 100 miles results in smaller standard errors.

overall decline. In the full draft we add in several different fixed effect combinations and time trends.

Table 1 - Baseline Estimates, Fertility 1910-1930

	(1)	(2)	(3)	(4)
Ag. Crop Index	-0.128**	-0.144**	-0.126**	-0.148***
	(0.055)	(0.057)	(0.050)	(0.052)
Population Restriction		Y		Y
Controls			Y	Y
Observations	33,374	29,960	32,141	28,727
Panel B: Y = #Children Under 5				
Average Ag. Crop Index	-0.171***	-0.245***	-0.240***	-0.079**
	(0.023)	(0.037)	(0.034)	(0.031)
Controls	Y	Y	Y	Y
Age Restriction		Y	Y	Y
Population Restriction			Y	Y
Same States as County Sample				Y
Observations	49,092,654	25,468,347	12,040,507	5,860,021

Notes: Panel A estimated using the county-level birth records. Panel B uses the complete count for 1910-1930. Every regression includes county and year fixed effects. Panel A controls include separate interactions between year fixed effects and baseline fractions of population who are non-white, between the ages of 6 and 14, and living in an urban area, as well as interactions with a baseline manufacturing output per capita. Panel B controls include dummy variables for age, race, marital status, farm status, and population size of place of residence. The population restriction drops the counties which are in the top 10 percent of population in 1910. The age restriction in Panel B is for women under the age of 35. Standard errors clustered at the state level: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The instantaneous decline in births could simply reflect retiming due to the war disruption. Indeed, a model that allows for household savings would predict that the number of children born should fall in the current period when female wages rise, but the increase in savings would result in more children in the future. We now turn to the complete count data samples in Panels B. In Panel B, we report the estimated impact of the average agricultural price index over the previous 5 years on the number of children under the age of 5 in the household. We estimate that a doubling of the price index over a 5-year period results in 0.171 fewer children under the age of five in the home. During our sample period, the average 5-year price index was 1.35 and the average number of children under age 5 was 0.61. Thus, on average, we estimate that the agricultural boom reduced the number of young children by about 9.8 percent relative to the mean. For younger women, who are more likely to be of childbearing age, we estimate a larger effect. For an average doubling of the index over 5 years, the coefficient increases to 0.245. Given slightly higher rates of young children in the home (0.85), the relative decline in young children among women under age 35 is 10 percent. For these same young women living in rural areas, the effect is similar, we find a 9.3 percent decline in children under age 5. Thus, in the intermediate run, there seems to be no evidence that the estimates from the county-level data reflect retiming alone.

One caveat to interpreting the estimates between panels A and B is that the complete count data include more states, and in particular, regions of the country where the agricultural boom was much more pronounced (i.e., the South). To ensure comparability between the complete count sample and the county sample estimates, we restrict the complete count sample to the states that appear in our annual county-level panel. We report these estimates in Panel B, Column (4). When the sample is restricted, we find that the coefficient remains negative (-0.079). Still, the smaller estimate implies a 2.5 percent reduction in the relative fertility rate when evaluating at the sample means. There are two likely explanations for the difference in coefficient magnitude. The first is that the counties in our restricted sample are those who experience less severe swings in crop prices. The second is that the labor intensity of agriculture is lower in the restricted sample, which may dampen the opportunity cost channel.

Our findings are strikingly similar to those reported by Schultz (1985), who examines the impacts of commodity price shocks to crops that utilize gender specific labor in Sweden (butter and grain). Our estimates are also similar in magnitude to those reported by Aaronson, Lange, and Mazumder (2014) for cohorts who were completely exposed to the rollout of Rosenwald schools (7.2-9.5 percent) during roughly the same era in the American South. The estimates are also similar to those reported by Bleakley and Lange (2009), who explore the impact of hookworm eradication in the American South. Our findings depart from that of recent work by Black, Kolesnikova, Sanders, and Taylor (2013), who document increases in fertility during the Appalachian Coal Boom. As we highlighted earlier, there is direct evidence that female agricultural wages respond to fluctuations in the agricultural price index, thus our findings are consistent with work by Schaller (2016) who shows that fertility is decreasing in the female wage.

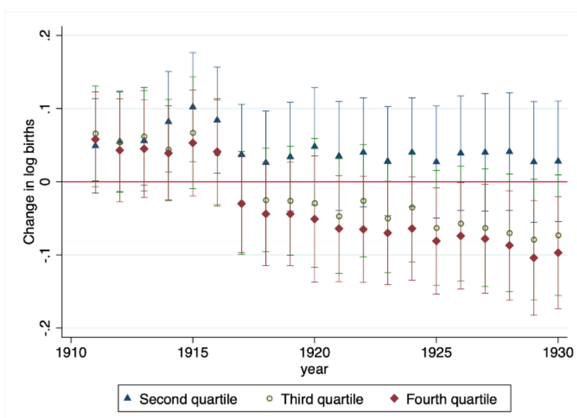
5.2 Quartile Event Study Results

The annual, county-level birth data make it possible to estimate the fertility response to crop prices over time. Instead of estimating a single coefficient on the price index, we modify equation 1 by interacting year fixed effects with separate dummy variables for the top three quartiles of the average price index at the WWI peak, 1916-1919:

$$LogBirths_{c,t} = \alpha + \sum_{k=2}^{4} \delta_{k,t}\, \tau_t \times 1(Index\ Quartile_c == k) + X'_{c,t}\beta + \tau_t + \sigma_c + \epsilon_{c,t}. \quad (3)$$

The $\delta_{k,t}$ coefficients on these interaction terms, which we plot in Figure 6, capture the deviation in births relative to the counties in the bottom quartile. The first thing to note is that each quartile had fairly similar birth trends leading up to WWI, providing a check on the assumption that underlying trends are not driving our main results. Once prices began to rise in response to WWI-induced demand, the birth patterns of each quartile diverge according to the intensity of their respective price boom.[11] While many of the 95% confidence intervals overlap, the pattern is consistent with both the expected timing and intensity of this natural experiment.

Figure 31 – Quartile event study results



Notes: Coefficients on the interaction between year fixed effects and dummy variables for the top three price index quartiles. 95% confidence intervals based on county-level clustering shown.

[11] The main results are similar when restricting the analysis window to end immediately after the post-WWI bust; the most substantial identifying variation is due WWI-related shocks.

5.4 Robustness – Other Potential Confounds

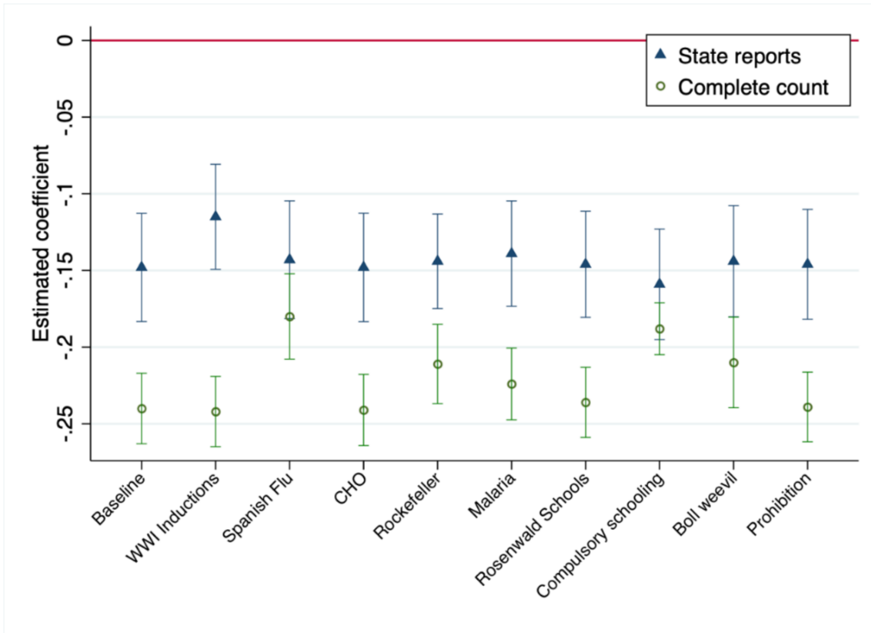
Given our sample period, there are several potential confounds that raise concern. First and foremost, our estimates could be the byproduct of a reduction in the availability of prime age males as a result of WWI. Furthermore, as men returned from Europe, America experienced the Spanish Influenza Pandemic (Almond, 2006; Beach, Ferrie, and Saavedra, 2018) which hit prime age males especially hard, which may have also affected fertility patterns. Our sample period also contains several secular movements that likely alter the returns to children, either through changes in the underlying health risk or through their future labor market returns. For example, the period we examine is squarely in the middle of the public health and high school movements. The arrival of boll weevil has also been shown to impact fertility. Finally, changes in alcohol consumption as a result of prohibition could also affect fertility.

In Figure 4, we present the estimates when controlling for these additional confounds. The first series, indicated by triangles, shows how these controls affect the coefficient on the crop index in regressions using the state report data on the annual number of births at the county-level. The second series, indicated by hollow circles, shows how the same control variables affect the estimates when using the complete count sample with the number of children under age 5 as the outcome of interest. Beginning on the left of Figure 4, we first replicate our preferred baseline estimates from Table 1. Moving from left to right, we add controls to account for the following potential confounders: the number of men drafted into WWI, exposure to the Spanish Influenza,[12] the rollout of County Health Organizations (CHO), Rockefeller Foundation's Hookworm eradication campaign, exposure to malaria, the presence of a Rosenwald School, changes in compulsory schooling laws, county-level boll weevil infestation, and county-level prohibition policies. It is clear from Figure 4 that the inclusion of these additional controls has little impact on our estimates of how the

[12] Note than the complete count sample is much smaller when we control for the impact of the Spanish Flu due to the availability of county-level mortality data.

agricultural index relates to either outcome of interest.[13] In the full version of the paper we also run several specifications to address concerns regarding migration.

Figure 4 – Robustness checks



Notes: Coefficients on the one year lagged crop index when using the annual county-level births as the outcome variable (state reports) and the 5-year average crop index when using the number of children under age 5 as the outcome variable (complete count) with the inclusion of different control variables. 95% confidence intervals based on state-level clustering show.

5.5 Long-Run Fertility Effects of Agricultural Boom

We now examine how variation in the price index over the prior 10 years affects the number of children in the household. In a static model, rising opportunity cost unambiguously decreases fertility along the intensive and extensive margins. In a dynamic model with additional margins of adjustment, the effects are less clear. Savings from the boom period may allow families to retime and replace or expand. To the extent that children are complements (substitutes) to the farm capital, the number of children may increase (decrease). Likewise, if farms are more profitable, or pay higher wages, it may influence the decisions of children to stay on the farm longer as they mature, as highlighted by Rosenzweig (1977). Therefore, a change in the number of children in the

[13] We include additional robustness checks in Appendix Table 3 showing that more general concerns about differential trends do not affect our county-level estimates.

household could reflect changes in when children exit the household rather than changes in the number of children born.

Measuring the long-run fertility effects with the complete count data recommends slightly different sample restrictions. We would ideally observe completed fertility or the total number of children ever born, but the closest variable collected in each of the Census years from 1910-1930 is the total number of children in the household at the time of enumeration. Younger women have not yet finished having children while older women will no longer have all of their children still living at home. For the long-run fertility regressions, we restrict our analysis to women between the ages of 30 and 40. We selected this age window because it will better capture the total fertility of women than the under 35 group. Furthermore, we limit the influence of endogenous household exit of children on our estimates by excluding women who are older than 40. The final two columns of Table 2 show that women between the ages of 30-40 have more total kids in the household but fewer children under the age of five than the under 35 group. They also have more children of any age than the full sample, reflecting the influence of household exit. While the summary statistics suggest that the total number of children for women aged 30-40 is the closest measure to completed fertility reported in the Census, children may have already exited the household at the time of enumeration and some women in the sample have not completed their fertility.

In Panel A of Table 2, we report estimates of how the average agricultural price index over the previous 10 years related to the total number of children in the home. Here we find that for a doubling in prices over a 10-year period, there is a statistically significant 0.427 reduction in total children relative to a sample mean of 2.19 children. Given an average 10-year price index of 1.19, this suggest that on average, there was a 3.7 percent reduction in children. Focusing on women who were between the ages of 30-40, we estimate a statistically significant negative coefficient of 0.654. Evaluated at the mean number of children and the average price index experienced by women aged 30-40, this implies a 4.8 percent decline in children. Further restricting to rural areas, we estimate a negative coefficient of 0.661, which translates to a 4.25 percent decline in total children. In Column (4), we report the estimate restricting the sample to include the same states used in the county-level annual birth sample. Here the coefficient remains negative but becomes much smaller in magnitude. Later in the paper we explore sources of heterogeneity, some of which may be related to this result. Finally, in Columns (5) and (6), we report estimates from samples that restrict to people living in their state of birth (Column 5), or to mothers born in the United States (Column 6). These restrictions have little impact on the estimated coefficient, alleviating concerns that our results reflect selective migration.

How could WWI and elevated agricultural prices lead to a permanent drop in fertility? As Galor (2012) and Aaronson, Lange, and Mazumdar (2014) note, increasing the price of child quantity decreases fertility along both the extensive and intensive margins. Thus, one possibility is that women had fewer children. Another is that more women remained childless. The next two panels of Table 5 explore the extensive and intensive margins of fertility. In Panel B of Table 2, we examine changes at the extensive margin, using an indicator that equals 1 if the woman has any children in the household, and is zero otherwise. In column 1, we report estimates using the sample of all women. We estimate a negative 0.069. Evaluated at the mean agricultural price index throughout the sample (1.19), this implies 1.29 percent decrease in the probability of a woman having children in the household. Focusing on agricultural households and women of prime age, the magnitude of the coefficient increases to -0.08. Evaluated at the index mean, this is a 1.5 percentage point reduction in the probability a woman has children in the household. The relative magnitude of this effect is quite large. For rural women aged 30-40, only 14 percent of women had no children in the household. Thus, elevated agricultural prices increase childlessness by about 10 percent relative to the mean. Consistent with Aaronson, Lange, and Mazumdar (2014), the fertility response is also driven by changes along the intensive margin. Using the same set of sample restrictions, Panel C shows that, conditional on having any children in the home, women had fewer children because of higher crop prices. To put this in context, the coefficients in Panel C which range from -0.328 to -0.506 are around two-thirds the size of the coefficients in Panel A.[14] The combined results of Table 2 indicate that both intensive and extensive margin responses meaningfully contributed to declines in overall fertility.

[14] The combined results from Column 4 suggest that the limited variation in crop prices for the state report sample primarily affected fertility through the extensive margin.

Table 2 - Number of Children in Home

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Y = # of children in home						
Average Ag. Crop Index	-0.427***	-0.654***	-0.661**	-0.035	-0.684***	-0.698**
	(0.143)	(0.181)	(0.220)	(0.147)	(0.241)	(0.219)
Panel B: Y = Any kids						
Average Ag. Crop Index	-0.069***	-0.078***	-0.080***	-0.043***	-0.078***	-0.086***
	(0.014)	(0.012)	(0.014)	(0.013)	(0.018)	(0.015)
Panel C: Y = (# of children in home | any kids > 0)						
Average Ag. Crop Index	-0.328**	-0.506***	-0.408*	0.117	-0.499**	-0.491**
	(0.147)	(0.183)	(0.203)	(0.135)	(0.216)	(0.202)
Age Restriction		Y	Y	Y	Y	Y
Population Restriction			Y	Y	Y	Y
Only States in County Sample				Y		
Live in State of Birth					Y	
Not Foreign Born						Y
Observations (A and B)	49,092,654	18,138,191	7,887,813	4,213,270	5,259,385	7,317,018
Observations (C)	37,846,641	14,835,267	6,768,714	3,607,568	4,544,434	6,265,822

Notes: 1910-1930 complete count census data. Each regression includes county and year fixed effects as well as dummy variables for age, race, marital status, farm status, and population size of place of residence. The population restriction drops the counties which are in the top 10 percent of population in 1910. The age restriction is for women between the ages of 30 and 40. Standard errors clustered at the state level: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

6. Conclusions

Identifying the causal relationship between income and fertility continues to be an important goal for economists, one with substantial implications for demography and policy design. In this paper, we presented evidence that income shocks that change the wages and opportunity costs of women can have significant impacts on fertility decisions, along both the extensive and intensive margins. The agricultural boom-and-bust of WWI, which differentially affected the wages and output prices of farms across the US based on their preexisting composition of crops, provided the necessary exogenous variation to study how families responded to income shocks. Analysis using newly-digitized annual birth records revealed that counties with larger price shocks experienced larger fertility declines in the short-run. This pattern was confirmed using an alternative measure of short-run fertility (children under the age of 5) in the complete count Census data. An important follow-on question is whether women simply retimed their fertility to take advantage of temporary wage increases. This does not appear to be the case: results focusing on the total number of children in the household from the complete count support the conclusion that these price shocks lowered the long-run fertility of affected women. These effects tended to be permanent, as a larger share of women never had children. Consistent with the notion that rising female wages (higher opportunity costs), were the dominant mechanism, we document that women delayed marriage in response to the elevated agricultural prices. In areas with higher rates of farm owner operators, the declines were partially offset due to the wealth effects as land values appreciated. There may have also been a limited role for investment in farm capital, yet these findings are less conclusive. Although data limitations prevent us from ruling out the importance of alternative factors, the bulk of the evidence points to female wages as being a significant contributor to the decline in fertility over this time period.

Much of the literature on fertility and family size focuses on the theoretical tradeoff between quality and quantity or how costs, both opportunity and explicit, have changed over time. Our work compliments this literature and shows that these are the primary underlying forces at play, even though children may directly affect family income in a rural setting. This additional dimension is still present in agricultural societies today even if many other factors (cultural, institutional, etc.) differ from those in the US in the early 20th century. Although the roles and expectations of children have changed substantially over the last century, parents still consider economic opportunities when deciding how many children to have. As policymakers consider potential responses to the many challenges that declining fertility presents to existing social programs, the results of this paper serve as a reminder that gender-specific changes to opportunity costs are central to the discussion. Put another way, there is more to the relationship between income and fertility than just income effects.

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