

Child Labor and the Labor Supply of Other Household Members: Evidence from 1920 America

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Although there is no lack of empirical evidence on the *determinants* of child labor (see Eric Edmonds and Nina Pavcnik, 2004; Alessandro Cigno and Furio C. Rosati, 2005, and references therein), much less is known about the *effects* of child labor. This paper tries to partially fill this gap by investigating the effect of an exogenous increase in a child's labor force participation on household labor supply decisions among urban households in early-twentieth-century America using micro data from the 1920 U.S. Census.

This paper is not the first to empirically investigate the effect of changes in one individual's labor supply on other household members' decision to work. Typically, these analyses refer to the effect of a husband's unemployment on his spouse's labor supply decision (Shelly Lundberg, 1985). Little is known, however, about the effect of children's changes in labor force participation on their parents' and siblings' labor supply, and this paper attempts specifically to address this issue. It is plausible that it is parents who decide on their children's use of time and retain control over their children's earnings. By analyzing how work effort is real-

located in a household in response to an exogenous increase in a child's labor market participation, one can learn something about parents' preferences and the constraints and opportunities they face in deciding on their children's time use.

One major difficulty that arises in identifying the household labor supply consequences of an increase in a child's work is that the labor supply decisions of all individuals in the household are likely to be taken simultaneously and to reflect both the same household characteristics and the same labor market conditions. This induces some spurious correlation between the labor supplies of different household members that makes it hard to ascertain how the household adjusts its labor supply as a consequence of one child entering the labor market. In order to circumvent this problem, one needs a source of variation in each child's labor force participation that is exogenous to the labor supply decisions of all other household members. In this paper, I suggest exploiting the variation in child labor induced by child labor laws. Because in 1920 different U.S. states had different legal minimum working ages, children of the same age happened to be eligible for work—i.e., at least as old as the minimum working age—in some states, but not in others. This allows us to derive a simple estimator of the effect of child labor laws on children's labor force participation based on the interaction of a child's age with his state of residence. To the extent that work eligibility affects a child's probability of work and, only via this channel, the labor supply of his household members, this can be used as an instrument for child labor.

Based on this strategy, I show that while becoming eligible for work makes a child significantly more likely to work and less likely to attend school than his siblings, this implies only a modest increase (fall) in the overall number of children working (attending school) in the household. I take this as evidence of a positive spillover in the labor supply of children living in

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the same household. Finally, I show that a rise in the proportion of working children in the household, once appropriately instrumented with the proportion of work-eligible children, has no statistically significant effect on parents' labor supply.

The structure of the paper is as follows. Section I introduces the data and presents some descriptive evidence on children's and parents' labor supply in 1920 America, and on the effect of child labor laws on children's labor force participation. Section II presents the empirical strategy. Section III presents the regression results. Section IV discusses the results and concludes.

I. Data and Descriptive Statistics

For the purpose of this analysis, I use micro data from the 1-percent sample of the Integrated Public Use Microdata Series (IPUMS) version of the 1920 U.S. Census (Steven Ruggles et al., 2004). To simplify my analysis, I focus on children age 10 to 16 living in urban areas.¹ The data provide information on usual employment and school attendance of all individuals.² One drawback of the data is that no information on work intensity (e.g., hours or days of work) is available. I integrate these data with information on minimum working age by state, as in Claudia Goldin and Lawrence F. Katz (2003). Following these authors, I compute two variables for each state: the age at which a youth can obtain a work permit for work during normal school hours, and the sum of school entrance age (in 1912), plus the education required to receive a work permit. I define the minimum

¹ One would expect child labor laws to have a stronger effect on urban children, since enforcement is probably more lax in rural areas and because many rural children have the opportunity of working on the household farm, where child labor laws generally do not apply or are hard to enforce. This complicates the analysis by inducing a further margin of adjustment in both children's and parents' labor supply.

² Work is defined based on the census variable "occ1950." A worker is everyone with an occupation code less than 980 (i.e., I exclude the following occupations: invalid/sick/disabled, keeps house/house work/housewife, at school, retired, inmate/prisoner, other nonoccupation, N/A). This includes work in the market and in the household enterprise. The Census also records whether the individual attended "school, college, or any educational institution at any time since September 1, 1919."

TABLE 1—CHILDREN'S LABOR FORCE STATUS BY AGE AND GENDER

Age	Boys		
	Work	School	> = Minimum working age
10	0.007	0.969	0.000
11	0.010	0.974	0.000
12	0.018	0.967	0.052
13	0.033	0.965	0.075
14	0.123	0.898	0.654
15	0.296	0.715	0.922
16	0.548	0.464	1.000
Total	0.139	0.859	0.362

Age	Girls		
	Work	School	> = Minimum working age
10	0.004	0.965	0.000
11	0.004	0.974	0.000
12	0.004	0.964	0.057
13	0.010	0.962	0.075
14	0.072	0.889	0.648
15	0.188	0.736	0.924
16	0.422	0.486	1.000
Total	0.097	0.858	0.372

Notes: The table reports the proportion of urban children in work and school by age and gender in 1920 America. The last column reports the proportion at or above minimum working age in their state of residence. Number of observations; 58,633. Source: IPUMS 1920 Population Census.

working age as the maximum of these two quantities.³

Table 1 reports some descriptive statistics on children's labor supply by age, separately for boys and girls. Participation rises sharply with age, although this rise is more pronounced for boys than for girls. While at age ten, less than 1 percent of both boys and girls are working, participation among 16-year-olds is 55 percent for boys and 42 percent for girls. Conversely, school enrollment declines rapidly with age, especially for boys. While around 3 percent of both boys and girls age ten are out of school, by age 16 this proportion rises, respectively, to 54 percent and 51 percent. Not all children in work are legally allowed to work and not all children who are legally allowed to work in fact work. In column 3, I report the

³ I drop individuals living in two states (Wyoming and New Mexico) for which no information on minimum working age can be derived from the data. Their exclusion is unlikely to affect substantially my results, since these individuals account for only about 0.5 percent of the overall sample.

proportion of children at each age who are work eligible. While no child younger than age 12 in the sample is allowed to work, by age 16 the corresponding proportion is 100 percent. Note that about 58 percent of children live in states with a minimum working age of 14 and around 27 percent live in states with a minimum working age of 15, with these two groups of states accounting for about 85 percent of the sample.

An analysis of parents' labor supply—not reported in the table—shows that virtually all fathers work, with participation on the order of 97 percent. Mothers' participation is rather low, about 12 percent.

In order to get a visual impression of the effect of legal minimum working age on children's labor force participation, I report the proportion of working children by age and gender in five groups of states (Figure 1), which I have classified based on their minimum working age (from 12 to 16). The graphs on the left side refer to boys, while the graphs on the right side refer to girls. A vertical line refers to the age just below minimum working age. To the right of this line, children are legally allowed to work. If labor laws bite—i.e., if children would work in their absence—and in the case of full compliance, one would expect participation to be zero before minimum working age and to be positive thereafter. If compliance is imperfect (but nonzero) some children will participate even before they turn legal working age, but one would still observe a rise in participation as a child becomes work eligible. Although one should be cautious in making inferences based on graphs referring to different groups of states, since these graphs are based on samples of different sizes and, hence, they vary substantially in their precision, the data appear roughly consistent with these predictions; participation stays roughly flat or rises modestly up to one year before a child turns legal working age, and then starts to increase.

Even in the absence of child labor laws, though, participation is likely to increase with age, perhaps because children's productivity rises and/or the disutility of labor falls as a child ages. This means that the age participation profiles in Figure 1 confound the effect of children's aging with the fact that they become work eligible. To try to isolate the effect of child labor laws, in Figure 2, I report the difference between the age participation profiles in each group of states relative to states with a

minimum working age equal to 14 that—as said—account for the greatest share of the sample. I standardize this difference to the difference in participation at age ten. If one postulates that, except for the effect of child labor laws, children's participation grows with age at the same rate across states, and if child labor laws have an effect, one would expect this difference to be zero for equally aged children who are either work ineligible or work eligible in both groups of states, and to be positive (negative) for children who are work eligible (ineligible) in their state but are work ineligible (eligible) in states with a minimum working age of 14. In Figure 2, I superimpose on the data two vertical lines in between which one would expect to see nonzero differences in participation. For example, in the bottom graph (MWA = 16), these two lines correspond to ages 13 and 16, since children age 14 to 15 should be less likely to work in states with a minimum working age of 16 compared to children in states with a minimum working age of 14. The data appear roughly consistent with the predictions of the simple model of child labor laws I have postulated. The differences in participation profiles are usually small up to the first vertical line; on the right of this line they become positive or negative—depending on which group of states I consider (minimum working age lower or higher than 14). After the second vertical line, they tend to converge back to zero. Differences appear particularly pronounced in states with a minimum working age equal to 15 and 16 that account for 27 percent and 8 percent of the sample, respectively. The results are less precise for states with a minimum working age equal to 12 or 13 that account for 5 percent and 2 percent of the overall sample, respectively. Finally, note that it also appears that this convergence is somewhat slower than one would expect based on the simple model postulated above. For example, at age 15 (16), children in states with a minimum working age of 15 (16) still participate less than those in states with a minimum working age of 14. On average, though, differences in participation between eligible and ineligible children appear consistent with the postulated effect of child labor laws. Also note that, except in states with a minimum working age of 12, the relative patterns of labor force participation are remarkably similar across gender groups.

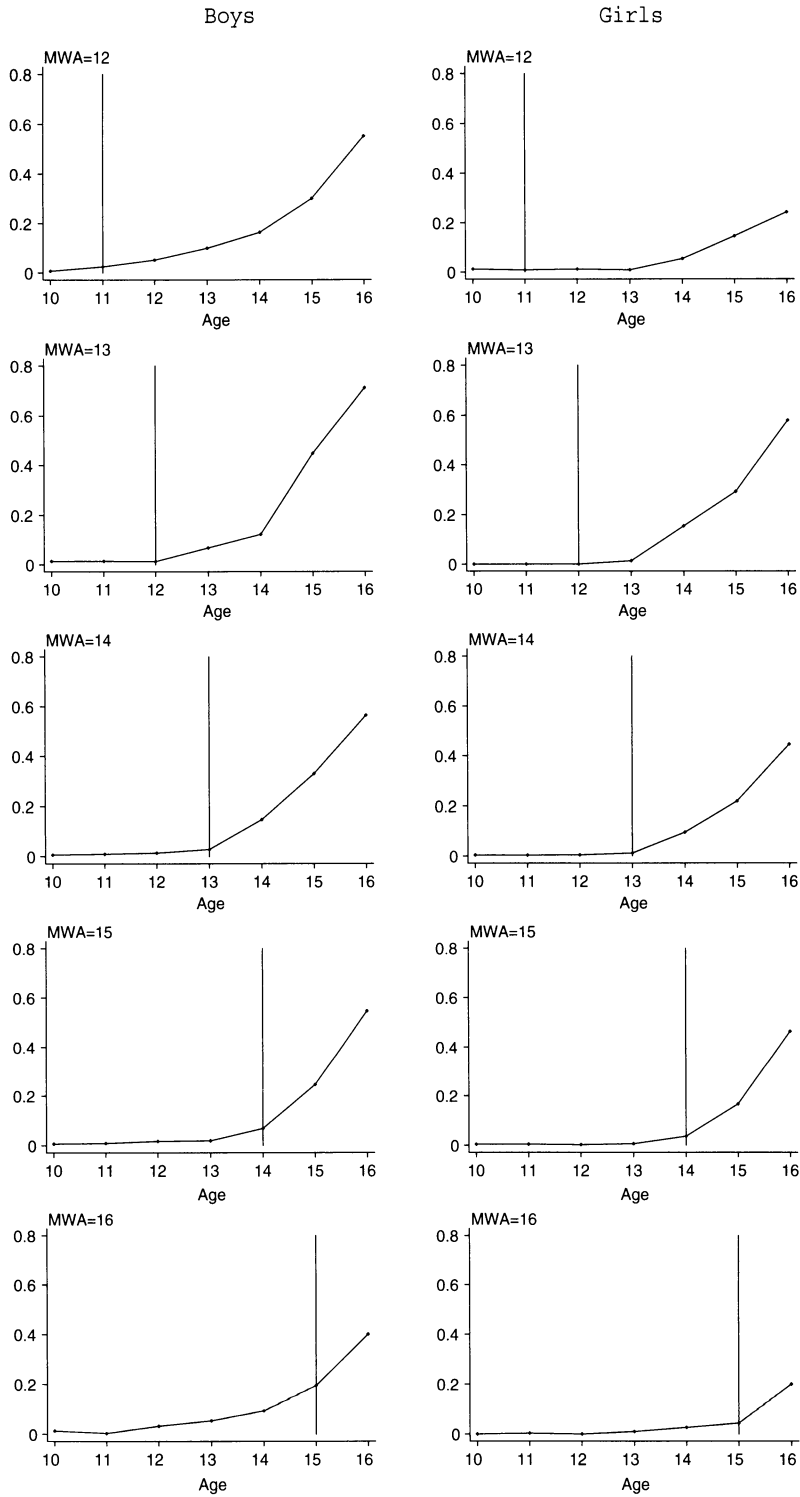


FIGURE 1. CHILD LABOR BY AGE AND MINIMUM WORKING AGE (MWA) IN THE STATE OF RESIDENCE

Notes: The figure reports the proportion of working children by age and sex in five groups of states, classified based on minimum working age (MWA). A vertical line refers to the maximum age at which children are not allowed to work in their state of residence (MWA-1).

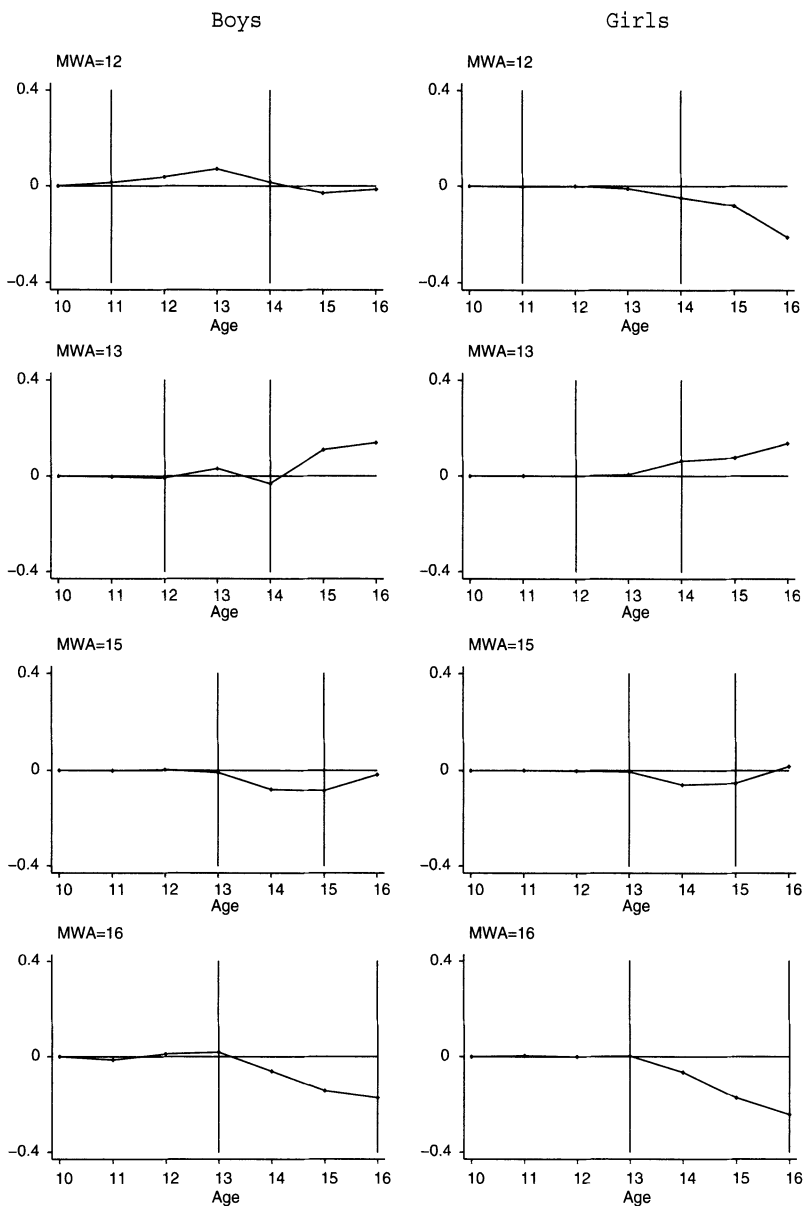


FIGURE 2. CHILD LABOR BY AGE AND MINIMUM WORKING AGE IN THE STATE OF RESIDENCE
*Differences (relative to states with minimum working age 14) in differences
 (relative to age 10)*

Notes: The figure reports the difference between the proportions of working children at each age in four groups of states defined based on their minimum working age (12, 13, 15, 16) relative to the proportion of working children in states with minimum working age 14. All series are standardized to the difference in participation at age ten. Two vertical lines refer to the interval where one would expect these differences to be nonzero (see text for details).

II. Empirical Strategy

In this section, I propose an empirical strategy aimed at assessing the impact of a child's work on his household's time-use patterns. I start by modeling the effect of a child's work eligibility on his own participation. Because households typically have more than one child, I then move on to separately estimating the effect of a child turning work eligible on this child's probability of work relative to that of his siblings and on the joint probability of work of this child and his siblings. I show that this provides a rather straightforward way of assessing whether any spillover in the labor supply of siblings exists. Last, I use the variation in the proportion of work-eligible children by household to identify possible labor supply responses of parents to their children's changes in participation.

Consistent with the discussion in the previous section, I postulate that, in the absence of child labor laws, each child's labor force participation grows with age at the same rate across states, and I model the effect of child labor laws as an additive term that increases participation similarly for all work-eligible children. Omitting other controls, this suggests estimating the following model:

$$(1) \quad T_{Asc} = \beta_0 + \beta_1 Z_{Asc} + f_{Ac} + f_{Sc} + u_{Asc},$$

where T_{Asc} is a 0/1 variable denoting whether child c of age A in state S is at work; the f_{Ac} 's are child's unrestricted age dummies common across states; the f_{Sc} 's are unrestricted state dummies common across all individuals in a certain state, irrespective of their age (hence, also common across all individuals in the same household); $Z_{Asc} = I(A_c \geq MWA_{Sc})$ is a dummy equal to one if a child's age is at least equal to the minimum working age—i.e., if the child is work eligible—in his state of residence; and u_{Asc} is an error term.

Identification in (1) is based on a simple differences (across children of different ages) in difference (across children living in states with different MWA) estimator that pools (the absolute value of) all the differences in Figure 2 and estimates the average gap in participation between work eligible and work ineligible children. Consistency of the OLS es-

timate of β_1 requires that a child's work eligibility is orthogonal to the error term in (1). I will provide suggestive evidence of this in the empirical section.

Model (1) ignores the potential effect of changes in a child's decision to work on his siblings' participation, which is ultimately one of the effects I am interested in estimating. Although in theory one might want to investigate how each child responds to each of his siblings' increased participation, in practice this exercise would require modeling a high number of cross effects. To cut through this problem, I simply concentrate on the effect of changes in the proportion of working children in the household on the probability of each child's work. Effectively, I augment equation (1) with the proportion of working children in household H denoted by T_{SH} . Because the proportion of working children in the household is itself a function of the children's age structure, and since this might have an independent effect on each child's decision to work, it seems appropriate to control additionally for the age structure of all children in the household. I do so by including the proportion of children in the household in each one-year age cell, d_{jSH} ($j = 10, \dots, 16$), and estimating the following regression:

$$(2) \quad T_{Asc} = \gamma_0 + \gamma_1 Z_{Asc} + \gamma_2 T_{SH} + \sum_j \gamma_{3j} d_{jSH} + f_{Ac} + f_{Sc} + v_{Asc},$$

where c is a child in household H . The coefficient γ_2 is a measure of the spillover in children's labor supply. This measures the effect of increased labor market participation of a child on his siblings' labor force participation. If $\gamma_2 < 0$ ($\gamma_2 > 0$), this is an indication of a positive (negative) labor supply spillover among children. Even in the absence of other sources of endogeneity, the OLS estimates of (2) will be inconsistent, since the variable T_{SH} on the right-hand side also contains the left-hand side variable T_{Asc} , inducing a classic division bias. To see how one can consistently estimate the coefficients in (2), notice that aggregating this equation across children in the household gives:

$$(3) \quad T_{SH} = \delta_0 + \delta_1 Z_{SH} + \sum_j \delta_{3j} d_{jSH} + f_{SH} + v_{SH},$$

where $\delta_1 = \gamma_1/(1 - \gamma_2)$ and Z_{SH} is the proportion of work-eligible children in the household. By replacing (2) into (3):

$$(4) \quad T_{Asc} = \gamma_0 + \gamma_1 Z_{Asc} + \gamma_2 \delta_1 Z_{SH} \\ + \sum_j \gamma'_3 d_{jSH} + f_{Ac} + f'_{Sc} + e_{Asc}.$$

Equations (2) to (4) are the basis of my empirical strategy. Equation (3) relates average participation among children in a household T_{SH} to their average work eligibility Z_{SH} . Equation (4) relates a child's work decision to his own and the average work eligibility, Z_{Asc} and Z_{SH} . To the extent that Z_{SH} is orthogonal to the error term in (2), one can simply use equation (3) as a first-stage equation for T_{SH} in (2) and estimate γ_1 and γ_2 via means of 2SLS (assuming $\gamma_2 \neq 1$ for identification). In this case, (4) has the natural interpretation of a reduced-form equation.

In order to understand the identification, note that since on the right-hand side of (4) I control for the household average of all the individual specific variables (age and work eligibility), the OLS estimate of γ_1 in (4)—or, similarly, the 2SLS of γ_1 in (2)—is simply the *within*-household estimator of a child's work eligibility. This is the estimated difference in participation between two siblings, one of whom is work eligible and the other who is not, once the difference in their age is accounted for. The OLS estimate of the coefficient on average eligibility in (3), δ_1 , instead is the *between*-household estimator of child eligibility. This coefficient picks up the effect of a child turning legal working age on the average number of working children in his household. Assume that $\gamma_1 > 0$ and $\gamma_2 < 1$. If $\gamma_2 < 0$ ($\gamma_2 > 0$), the between-household estimator is below (above) the within-household estimator, in which case the effect of a child turning work eligible on the average probability that a child in his household moves into work is a fraction (a multiple) of the positive effect that turning work eligible has on this child's probability of work. This means that this child's siblings will be less (more) likely to work. I interpret the difference between these two coefficients as being due to the positive (negative) spillover in children's labor supply.

Models (2) to (4) suggest that one can estimate the coefficients γ_1 and γ_2 (and δ_1 , which is a nonlinear combination of the two) via a 2SLS estimator, where the proportion of

working children in a household is instrumented by the proportion of work-eligible children. Consistency of the 2SLS estimator requires each child's work eligibility to be uncorrelated with the unobserved determinants of both this child's and his siblings' labor supply. Effectively, child labor laws are a good instrument for child c 's labor if child c 's work eligibility affects his siblings' work only via changes in child c 's labor. Again, in the empirical section, I will provide some circumstantial evidence on the validity of the exclusion restriction.

As a last step, I examine parents' labor supply responses to changes in child labor. To do so I regress parents' labor force participation on children's average work participation T_{SH} , the children's age structure d_{jSH} , and state fixed effects b_{SH} :

$$(5) \quad T_{SHP} = \phi_0 + \phi_1 T_{SH} + \sum_j \phi_3 d_{jSH} \\ + b_{SH} + w_{SHP},$$

where P denotes a generic parent ($P = F, M$ for father and mother, respectively). The coefficient ϕ_1 picks up the effect of changes in the proportion of working children on parents' labor supply. There are good reasons to be skeptical about the consistency of the OLS estimate of ϕ_1 in (5), since unobserved determinants of parents' work, whether on the side of labor demand or supply, are likely to affect simultaneously children's labor supply, hence, inducing a bias in the OLS estimate of ϕ_1 . The direction of the bias is hard to determine *ex ante*. For example, low household income or higher demand for labor in the households' area of residence might make both parents and children more likely to work, leading to an overestimate of ϕ_1 . In contrast, if exploitive parents tend to work less while their children tend to work more, or if a fall in parents' wages makes parents less likely to work and children more likely to work (via an income or substitution effect), this will lead to a downward bias in the OLS estimate of ϕ_1 . If, however, Z_{SH} is orthogonal to w_{SHP} —i.e., if conditional on children's age structure and the household's state of residence, children's work eligibility is uncorrelated with unobserved determinants of parents' labor supply—one can obtain a consistent estimate of ϕ_1 by

TABLE 2—CHILD LABOR LAWS AND CHILDREN'S LABOR SUPPLY

	Coefficient on eligibility	Dependent variable		Sample	Children 10–16 age structure	Additional controls	Observations
		Work	School				
1. Individual (OLS)	Own	0.057*** (0.007)	−0.056*** (0.008)	All	No	No	58,633
2. Individual (OLS)	Own	0.056*** (0.007)	−0.057*** (0.008)	Restricted	No	No	54,392
3. Individual (OLS)	Own	0.055*** (0.006)	−0.056*** (0.007)	Restricted	No	Yes	54,392
4. Reduced form (Within)	Own	0.078*** (0.009)	−0.080*** (0.009)	Restricted	Yes	No	54,392
	Average	−0.043*** (0.014)	0.042*** (0.016)				
5. Reduced form (Within)	Own	0.077*** (0.009)	−0.079*** (0.009)	Restricted	Yes	Yes	54,392
	Average	−0.042*** (0.013)	0.042*** (0.015)				
6. First stage (Between)	Average	0.036*** (0.009)	−0.037*** (0.012)	Restricted	Yes	Yes	54,392
7. 2SLS	Average	−1.154* (0.616)	−1.118 (0.723)	Restricted	Yes	Yes	54,392

Notes: The table reports the effect of work eligibility on children's labor supply. Rows 1 to 3 report OLS estimates of equation (1) in the text; rows 4 and 5 of equation (4); rows 6 of equation (3); and row 7 the 2SLS estimate of equation (2). Data refer to children age 10 to 16. Restricted sample includes only children living with parents age 25 to 60. All regressions control for state fixed effects, child's age dummies interacted with a gender dummy, and four race dummies. Rows 3 to 6 also control for the proportion of children in the household age 10 to 16 in each one-year age \times sex cell. Additional controls include: dummies for total number of children age 10 to 16, dummies for the number of children age 0 to 9 in each one year age cell, dummies for household size, three dummies for residential status (not in metro area; in metro area—central city; in metro area—outside central city), three dummies for home ownership (owned and not mortgaged; owned and mortgaged; rented), father's and mother's race, dummies for whether the father and the mother are born abroad, dummies for whether the father and mother are illiterate, father's and mother's age and age squared, dummies for missing father and mother, and number of elder siblings. Clustered standard errors (by state and household children's age structure) in parentheses.

* Significant at 10-percent level.

** Significant at 5-percent level.

*** Significant at 1-percent level.

instrumenting T_{SH} with Z_{SH} , using equation (3). Effectively, the model conditions on the children's age structure and state fixed effects and exploits the interaction between these two variables for identification. Consistency of the 2SLS estimates requires differences (across households) in differences (across states) in children's age structure to be uncorrelated with unobserved determinants of parents' labor supply. Again, since the model is just identified, there is no direct test of the exclusion restriction. As is common practice though, I check for the robustness of the regression results to the inclusion of a large array of observable controls.

III. Regression Results

In this section, I present the regression results based on the empirical strategy just outlined. As

a first step, I estimate equation (1), which pools all individuals, whether from the same or different households. To start with, I include all individuals age 10 to 16 in the data. I report the estimates on work eligibility in Table 2, row 1. Since Figure 2 shows similar effects of child labor laws by gender, in this and all the other regressions in the Table, I pool boys and girls. As discussed, the model includes age and state dummies. In order to allow for boys' and girls' participation profiles to differ over their life cycle, I also allow the age dummies to vary unrestricted by gender. Finally, I add four dummies for child's race (white, Spanish, black and mulatto, and other). Standard errors in this and all the other regressions are clustered into groups based on the age structure of children in the household and state of residence. The regression delivers an estimated effect of child

labor laws (β_1) of 0.057 (s.e., 0.007), suggesting that turning work eligible makes a child more likely to work by almost 6 percentage points. To my knowledge, this is the first paper to show a significant impact of child labor laws on child labor in early twentieth century America.⁴

As a next step, I try to investigate the presence of spillovers in the household labor supply. To do so, I first need to identify the cohabiting parents and siblings for each child in the sample. I use the imputed location of the mother and the father in the household that is available in the IPUMS version of the 1920 Census to identify the children's parents, and I classify as siblings those who share the same parents. For this reason, I examine only children living with parents. To limit the possible misclassification arising from the imputation of the father and the mother, I restrict the sample to individuals with parents in the age range 25 to 60. This delivers a restricted sample of 54,392 children in 34,118 households (out of an original sample of 58,633 children in 37,579 households).

Using only children living with parents, one might be worried about the sample selection issue that potentially arises if—conditional on work eligibility—the probability of living with parents is correlated with the probability of work. To see this, consider the extreme case in which all work-eligible children in work do not live with their parents. In this case, the coefficient on the work eligibility variable will be zero, although, effectively, child labor laws tend to increase participation. Although this is unlikely to be a major problem for children in the

age range 10 to 16, who largely live with their parents, I try to test for this in row 2 of Table 2, where I report results for exactly the same specification as in row 1, estimated on the restricted sample of children (those living with parents age 25 to 60). The results are virtually unchanged from row 1 (coefficient is 0.056, s.e., 0.007), suggesting that children's endogenous living arrangements are unlikely to be a major source of bias in my estimates.

It is sometimes claimed that child labor laws might be endogenous to the local demand or supply for child labor.⁵ If it happens, for example, that child labor laws are less stringent in states where the latent level of child labor is higher, one will spuriously attribute the effect of market forces to child labor laws, hence overestimating their impact. To the extent that such higher latent levels of child labor are reflected in the labor supply of all children living in a certain state, irrespective of their age, model (1) accounts for this potential source of endogeneity by explicitly controlling for state fixed effects. As an additional check for this potential source of bias, I include a number of controls in the model. One would expect differential preferences for child labor regulation across states to reflect, at least in part, households' constraints and opportunities. So, if conditional on state and age fixed effects, such characteristics are correlated with child labor laws, one would expect their inclusion to affect the OLS estimate of β_1 in (1). Similar sources of bias might be induced by some parents endogenously adjusting their residential location or fertility decisions to the level of child labor laws. For example, if households that want their children to work but that originally reside in states with a high minimum working age relocate to states with a low minimum working age to take advantage of their children's work eligibility in the destination state, this might again induce some upward bias in the estimate of β_1 in (1). As an admittedly imperfect test for these potential sources of correlation, in row 3, I run equation (1) on the restricted sample (as in row 2) with a large array of controls. For each child, I include the number of elder (cohabiting) siblings. As household controls, I include the total

⁴ While there is evidence that compulsory schooling and child labor laws in force in early twentieth century America had a significant impact on children's school enrollment (Robert Margo and Aldrich Finegan, 1996; Goldin and Katz, 2003) and their subsequent educational attainment (see for example, Joshua D. Angrist and Alan B. Krueger, 1991; Daron Acemoglu and Angrist, 2001; Adriana Lleras-Muney, 2002), the empirical evidence on the effect of labor market legislation on labor market participation of children is much more scant. Carolyn M. Moehling (1999) finds that child (manufacturing) labor laws had little effect on children's employment between 1900 and 1910. Between 1910 and 1920, however, a strong movement advocating more stringent child labor legislation arose, and by 1920 all U.S. states had some form of child labor legislation. (For a fascinating account of the history of child labor legislation in the United States see Walter Trattner, 1970.) Differences in timing and data can potentially explain the difference between my results and those of Moehling (1999).

⁵ See Matthias Doepke and Fabrizio Zilibotti (2005) for a theory of endogenous adoption of child labor laws.

number of children age 10 to 16, the number of children age 0 to 9 in each one-year age cell, household size, residential status, and home ownership. As parents' controls, I include father's and mother's race, literacy status, nativity status, age and age squared, and dummies for missing father and mother. Finally, I control for the age and sex structure of children age 10 to 16 in the household by adding the proportion of children age 10 to 16 in each one-year age \times sex cell. The coefficient on work eligibility is very close to the one in row 2, (0.055, s.e., 0.006), implying no correlation between observed characteristics and child labor eligibility. I take this as an indication of work eligibility being unlikely to be correlated with unobserved determinants of child labor.

Row 4 of Table 2 tests for the presence of spillovers in the labor supply decisions of children. I start by estimating model (4), where, in addition to each child's own eligibility, I also include the proportion of work-eligible children in the household. This and the following models are estimated on the restricted sample. I start by estimating the model without any additional controls, but I still condition for each child's individual characteristics (age and sex) and, as already discussed in Section II, I include the average age (and sex) structure of children age 10 to 16 in the household. The estimated coefficient on "own eligibility" (γ_1) is 0.078 (s.e., 0.009). The coefficient on "average eligibility" ($\delta_1\gamma_2$) is -0.043 (s.e., 0.014). Recall that the first coefficient is the *within*-household estimate of the effect of child labor laws, while the second is the difference between the *between*-households and the *within*-household estimates. If the exclusion restriction holds, one can interpret the second coefficient as some measure of the spillover. As a partial check for the validity of my identification assumption, again, I reestimate equation (4) with the inclusion of additional controls (as in row 3). The estimated coefficients in row 5 essentially are unaffected by the inclusion of these controls lending some credibility to the exclusion restriction.

Rows 4 and 5 suggest that some positive spillover is at work. This can be seen equivalently by estimating equation (3), which analyzes the effect of increased eligibility on increased participation across households. Row 6 of Table 2 reports the effect of increased average eligibility on average participation (δ_1).

I report results with the entire set of additional controls, since I have shown that their inclusion leaves the coefficients of interest unchanged. The estimated effect is the sum of own eligibility (γ_1) plus the effect of average eligibility ($\delta_1\gamma_2$) in row 5. The estimated coefficient equals 0.036 (0.009). This suggests that one extra work-eligible child in the household is associated with a rise in the probability of work of either this child or his siblings of 3.6 percentage points. If one compares this result to the one in row 5, this suggests that about 55 percent ($= 1 - 0.036/0.077$) of the differences in eligibility across households do not show up in differences in average participation, as one might have inferred from looking at two children in the same household. I attribute this difference to a positive spillover in children's labor supply. Note that, as expected, the pooled estimate in row 3 is approximately half way between the *within*-household estimate in row 5 and the *between*-household estimate in row 6.

In row 7, I report the 2SLS estimates of equation (2). As in row 6, I report only results with the whole set of controls. I report only the coefficient on average children's work participation (γ_2), since the estimated coefficient on own eligibility (γ_1) is identical to the one from equation (4) reported in row 5. The estimated coefficient is -1.154 (s.e., 0.616). Note that this is the ratio of the coefficients on average eligibility in rows 5 (-0.042) and 6 (0.036).

Finally, column 2 of Table 2 reports the effect of child labor laws on children's schooling. It is interesting to note that child labor laws seem to have a pronounced negative effect on school enrollment which is almost the mirror image of the effect on employment. As shown in rows 5 and 6 of Table 2, an increase in the proportion of work-eligible children in the household—or the proportion of children allowed to drop out of school which seems to be the same—tends to decrease the probability that a child in the household works *and* tends to decrease the probability that this child drops out of school, although the 2SLS estimate in row 7 is not significant at conventional levels.

I have performed several robustness checks on my data (not reported). First, I have replicated my exercise only on children age 10 to 15. The results are invariant to this selection rule. Second, I have run separate regressions by

TABLE 3—CHILDREN'S WORK AND PARENTS' LABOR SUPPLY

Independent variable	Father	Father	Mother	Mother
1. Proportion working (OLS)	0.002 (0.004)	0.003 (0.004)	0.098*** (0.010)	0.040*** (0.008)
2. Proportion eligible	-0.008 (0.005)	-0.007 (0.005)	0.010 (0.012)	0.007 (0.010)
3. Proportion working (2SLS)	-0.199 (0.152)	-0.189 (0.150)	0.274 (0.355)	0.189 (0.283)
Observations	29,952	29,952	32,684	32,684
Controls	No	Yes	No	Yes

Notes: The table reports the effect of children's work (eligibility) on parent's labor supply. Row 1 reports OLS estimates of the coefficient on the proportion of working children (equation (5)). Row 2 reports the coefficient on the proportion of work-eligible children. Row 3 reports 2SLS estimates of the coefficient on the proportion of working children, using the proportion on work-eligible children as an instrument (equation (3)). All regressions control for state fixed effects and the proportion of children in the household age 10 to 16 in each one-year age \times sex cell. Additional controls include: total number of children age 10 to 16, the number of children age 0 to 9 in each one-year age cell, household size, three dummies for residential status (not in metro area; in metro area—central city; in metro area—outside central city), three dummies for home ownership (owned and not mortgaged; owned and mortgaged; rented), father's and mother's race, dummies for whether the father and the mother are born abroad, dummies for whether the father and mother are illiterate, father's and mother's age and age squared. Clustered standard errors (by state and household children's age structure) in parentheses.

* Significant at 10-percent level.

** Significant at 5-percent level.

*** Significant at 1-percent level.

number of children. Results are similar with evidence of a positive spillover at all parities, but individual coefficients are never statistically significant. This should be no surprise since these estimates are obtained on relatively smaller samples than the one used in Table 2.

I now turn to the effect of child labor on the labor supply of parents, based on equation (5). In the first row of Table 3, I report OLS estimates of the coefficient ϕ_1 in equation (5). The second row reports reduced-form estimates where parents' employment is regressed on the proportion of work-eligible children Z_{SH} plus the same controls as in row 1. Finally, in the last row I report 2SLS estimates where the proportion of working children T_{SH} is instrumented by the proportion of work-eligible children Z_{SH} , according to equation (3). The regression results for fathers are reported in the first two columns, while results for mothers are reported in the following two columns. For each parent, I report regressions with and without additional controls. All regressions, however, control for state fixed effects and the age and sex structure of children in the household age 10 to 16. Estimates refer to observations in the restricted

sample for which information on the father or the mother is available.

When using OLS, I find a small, positive, and insignificant correlation between fathers' and children's labor. When controls are included, the estimated coefficient is 0.003 (s.e., 0.004). The correlation between children's and mother's labor supply is higher and statistically significant: the estimated coefficient is 0.040 (s.e., 0.008) when all the controls are included, implying that one extra child in work in a household with three children is associated with a rise in the probability of mother's work of around 1.3 percentage points. As said, these coefficients are unlikely to provide an indication of parents' responses to children's changes in participation. For this reason, in the second row I report the coefficient on children's average work eligibility, i.e., the OLS estimate of the reduced-form equation. The coefficients are small: -0.008 (s.e., 0.005) for fathers and 0.010 (s.e., 0.012) for mothers. No substantial differences arise when controls are included, suggesting that children's work eligibility is likely to be orthogonal to the error term in (5). The coefficients are not statistically distinct from zero, so

one cannot conclude that there is evidence of an effect of children's increased work eligibility on parents' labor supply.

Finally, in row 3 I report the 2SLS estimates of (5) where I use (3) as a first stage equation for children's labor. One should be cautious in interpreting these coefficients because child eligibility affects not only the proportion of working children but also the proportion of children attending school. To the extent that these different margins of children's time use also affect parents' labor supply, the 2SLS coefficient of children's labor market participation also absorbs these other effects. It is still useful, though, to derive the 2SLS estimator in order to get a sense of the magnitude of the implied effects. When controls are included, the 2SLS estimate of the effect of children's labor market participation on father's participation equals -0.189 (s.e., 0.150). This is simply the ratio of the reduced-form estimate in Table 3, row 2 (-0.007) and the first stage estimate in Table 2, row 6 (0.036). For mothers, I find a positive effect of 0.189 (s.e., 0.283). These are rather large coefficients, implying that a 33-percentage-point rise in children's labor market participation reduces father's labor market participation by around 6 percentage points and increases mother's participation by the same amount, suggesting perhaps some complex patterns of substitution between mother's and father's labor supply as children turn work eligible. Since the reduced-form estimates are insignificant at conventional levels, however, it is not surprising to find that the 2SLS estimates are also imprecisely estimated and are not distinct from zero.

IV. Discussion and Conclusions

A useful way of thinking about a child turning work eligible is that this corresponds to an improvement in this child's labor market opportunities, i.e., a rise in his shadow wage. The first effect one would expect is a rise in this child's participation and a fall in schooling, as the opportunity cost of not working and being in school rises. There are also, however, potential effects on the labor supplies of other household members.

If the household pools resources, utility is additively separable in the leisure of different household members, and the budget constraint

is additively linear in these members' time use, then a rise in an individual's wage that is sufficiently high to push this individual into the labor market will potentially have an income effect on other individuals' labor supply. To the extent that leisure (schooling) is a normal good, this might imply a withdrawal of other household members from the labor market. This effect is reinforced if there are patterns of substitution in the utility function (i.e., different individuals' consumption of leisure or schooling are substitutes) or in the budget constraint (e.g., if individuals are substitutes in home production). Obviously, the opposite will happen if there are complementarities between different members' leisure in the utility function or other nonlinearities in the budget constraint (e.g., if there are economies of scale arising from more than one child attending school or going to work).

Theory provides little guidance on which of these effects will prevail, and this is ultimately an empirical matter. My estimates are unable to disentangle these different channels, but they can at least provide an indication of their combined contribution.

The results of the paper are threefold. First, based on a simple differences-in-differences estimator between children of different ages living in states with different minimum working ages, I find a positive significant effect of a child turning work eligible on his probability of work, and a negative effect on his probability of school attendance relative to his siblings.

Second, I investigate how differences in work eligibility affect the employment choices of children living in different households. I show that as the proportion of work-eligible children in a household rises, the overall participation of working children in that household rises too. I also find, however, that this effect is only a fraction of the effect that one would have inferred from comparing eligible and ineligible children in the same household. This suggests that as a child becomes work eligible, this not only increases his probability of work but also simultaneously reduces his siblings' probability of work and increases their probability of being in school. There is a positive spillover in children's labor supply.

Third, I investigate the labor supply responses of parents to children's increased work (eligibility). The 2SLS leads to rather imprecise

estimates and one cannot conclude that these effects are statistically significant.

Because of data limitations, I am unable to say whether any significant response at the intensive margin exists for parents or siblings. One has to bear this caveat in mind in interpreting the regression results.

If, as it seems reasonable, parents decide on their children's use of time and retain control over their children's earnings,⁶ these results suggest that *prima facie* it is not in parents' interests to withdraw from the labor market in response to a child's increased labor market participation. This might be due either to parental altruism (as in Sonia Bhalotra, 2004) or even to self-interest if parents derive some private utility from their children's increased human capital accumulation. This result is consistent with the "luxury axiom" (Kaushik Basu and Pham H. Van, 1998), according to which child labor is largely the result of extreme poverty (possibly coupled with poor access to credit, as in Jean Marie Baland and James Robinson, 2000), rather than parents' attempt to avoid working themselves. If, as is also plausible, children would rather have their parents work than work themselves, this result seems to rule out a misalignment between parents' and children's interests in connection to child labor, a theoretical possibility and one that would raise serious policy concerns (Christopher Udry, 2004).

By the same token, the empirical results suggest that it is in parents' interest to shelter some children from work as to encourage others to enter the labor market. Again, it is not possible, based on the data in this paper, to determine if this happens because of parents favoring certain children (either out of self-interest or altruism) or possibly because of the existence of patterns of substitution between the time use of different children. Arguably, the most novel result that emerges from this analysis is that some children in the household appear to work to the benefit—in terms of increased schooling and lower labor supply—of others. This obviously raises serious concerns about the distribution of work and schooling among siblings and about the desirability of child labor laws. Although child

labor laws seem effectively to work to a child's benefit, by raising a child's school attendance and lowering his labor at early ages, the cost of this seems to be borne partly by his siblings.

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⁶ See Donald O. Parsons and Goldin (1989) for evidence of working children not retaining control over their labor earnings.

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