Spillover Effects from Minimum Wages in Agriculture

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Abstract

The Adverse Effect Wage Rate is a minimum wage that must be paid to non-immigrant agricultural guest workers working in the United States under the H-2A visa program. The AEWR was established as a mechanism to prevent domestic farmworker wage depression resulting from an increase in the employment of foreign workers, but growers argue that the AEWR influences the wages of all other workers and that it is generally too high. In this paper, we provide empirical estimates of the effects of changes in the AEWR on the wages and employment of domestic farmworkers. We estimate fixed-effects panel regression models using a lagged AEWR variable as an instrument and a labor demand proxy variable (a Bartik instrument) to control for unobserved agricultural labor demand shocks. Our econometric analysis indicates that higher AEWRs cause the wages and employment of domestic farmworkers to increase. We find an elasticity of domestic wages (hours worked per employee) with respect to the AEWR of 0.54 (2.07). Our results indicate that a policy that would freeze the AEWR for one year would reduce the wage growth of domestic employees by $570 million.

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Introduction

The Adverse Effect Wage Rate (AEWR) is a state-level minimum wage that must be paid to foreign agricultural guest workers working under the H-2A visa program and the corresponding US farmworkers who work for H-2A employers.[1] The AEWR was originally implemented to help prevent domestic farmworkers from facing downward wage pressure as a result of competition from foreign workers, many of whom have low reservation wages due to the relatively poor economic conditions in their countries of origin (UFW v. DOL, 2020; Congressional Research Service, 2008). However, unlike other minimum wages, the AEWR is endogenously determined by market conditions, and there is general concern that it is too high (Critterden, 2020; Lewison, 2021). Growers who pay wages lower than their state's AEWR have reason to be concerned about a market-based minimum wage because the AEWR may influence the wages of domestic workers, causing the market wage to inadvertently rise.[2] In this paper, we quantify the extent to which changes in the AEWR affect the wages and employment of domestic farmworkers. We conclude by quantifying the labor-market impacts of using the AEWR as a policy tool.

In November of 2020, the United States Department of Labor (DOL) announced a rule that would have frozen the AEWR for two years and made changes to the calculation method. Even though the AEWR acts as a wage floor for foreign workers, some argue that it could serve as a wage ceiling for domestic workers because employers who advertise employment opportunities to domestic workers at the AEWR can simply recruit H-2A workers if domestic workers are unwilling to perform farm work at that wage rate (Congressional Research Service, 2008). Farm employer advocates claim that the AEWR operates as a de facto minimum wage for all agricultural workers and that changes to the AEWR methodology are

[1] Throughout this article we define US farmworkers as native-born and foreign-born farmworkers who are not working under the H-2A visa program and who are not working for an employer who employs H-2A workers. We refer to these workers as "domestic" farmworkers despite the fact that some of them are foreign-born and are present in the United States without legal authorization to work.

[2] The earnings of H-2A workers in a given year are included in the calculation of the AEWR for the following year. To the extent that H-2A workers are paid wages above the average wage for a given local labor market, the AEWR will overstate the true average wage of domestic workers in that market.

necessary to keep farmers profitable in the US (Critterden, 2020; Lewison, 2021). Regardless of whether the AEWR serves as a wage floor or ceiling for domestic workers, if the AEWR influences domestic farm labor market outcomes, changes to the AEWR will affect the wages paid to domestic farmworkers.

Due to shortages of domestic workers, the H-2A program has grown rapidly over the past decade (Castillo et al., 2021; Martin and Rutledge, 2022; Castillo, Martin, and Rutledge, 2022). In order to utilize the H-2A program, farm employers must provide evidence that "there are not sufficient able, willing, and qualified US workers available to perform the temporary and seasonal agricultural employment for which non-immigrant foreign workers are being requested" (DOL, 2021)[3] An emergent literature documents the rapid expansion of the H-2A program (Castillo et al., 2021; Castillo, Martin, and Rutledge, 2022), yet there is a gap in the literature with respect to evaluating the impacts of the AEWR.

Some argue that minimum wages cause unemployment for low-skilled workers, while others argue that the minimum wage is essential to prevent exploitation and improve the purchasing power of consumers (Card and Krueger, 1994, 2000; Neumark and Wascher, 2008). A number of studies have found evidence consistent with perfectly competitive labor markets in which raising the minimum wage above the existing wage reduces employment.[4] However, other findings suggest that higher minimum wages lead to higher employment (e.g., Card and Krueger, 1994; Machin and Manning, 1994; Dickens et al., 1994; Richards, 2018). More recenly, Cengiz et al. (2019) finds that minimum wages have no effect on employment and that the distribution of wages is simply shifted upwards to the new minimum. Moreover,

[3]One of the factors contributing to the rise in H-2A use is the diminishing supply of domestic farmworkers due to the expanding Mexican economy, lower birthrates among rural Mexicans, more off-farm employment opportunities for the rural Mexican population, and increased immigration enforcement (Taylor, Charlton, and Yúnez-Naude, 2012; Charlton and Taylor, 2016, 2020; Bampasidou and Salassi, 2019; Zahniser et al., 2018; Zahniser, Hertz, and Charlton, 2019; Kostandini, Mykerezi, and Escalante, 2013; Ifft and Jodlowski, 2016; Charlton and Kostandini, 2020). Domestic farm labor shortages, and thus the demand for H-2A workers, are also influenced by the decline in follow-the-crop migration (Fan et al., 2015). As fewer workers travel to work on farms, the geographic range of local labor markets diminishes, making labor shortages more prevalent during periods of peak labor demand (Fisher and Knutson, 2012).

[4]For instance, Kim and Taylor (1995) find negative employment effects in the retail trade sector, Singell and Terborg (2007) for eating and drinking workers, and Kandilov and Kandilov (2020) in agricultural sector, to name a few.

Ashenfelter and Jurajda (2022) find spillover effects on workers earning wages above the minimum wage such that the wage premium remains the same before and after the increase in the minimum wage. Fan and Pena (2019) discuss the potential for a "lighthouse effect" such that workers who are not directly covered by the minimum wage are influenced by it because it serves as a general signal about labor market conditions. Consequently, the effects of the AEWR on wages and agricultural employment is ambiguous and is left as an empirical question.

In a typical setting, minimum wages are exogenously determined by policymakers such that variation in the minimum wage variable can be directly used to identify minimum wage effects on wages or employment. However, in our setting, the AEWR for a given year is determined by the USDA's Farm Labor Survey from the previous year, which provides a measure of average wages from 17 different regions across the US. As as result, the AEWR is linked to recent market conditions, which creates an identification challenge. Thus, additional measures need to be taken to identify the causal effect of the AEWR on domestic farmworker labor market outcomes.

Our empirical model regresses domestic farmworker wages (or employment) on the AEWR using a fixed-effects panel model at the state-year level of aggregation. Our main identification challenge results from unobserved labor market shocks in the previous period that influence the AEWR. We derive expressions for two plausible sources of omitted variables bias (unobserved lagged labor supply and demand shocks) in Section 3. We address this issue using a two-pronged approach.

First, we develop a proxy for lagged agricultural sector labor demand shocks using a Bartik instrument and include it as a control variable.[5] As expected, our empirical results reveal that the inclusion of the Bartik control helps mitigate upward bias in both the employment

[5] A Bartik (or shift-share) instrument is named after Bartik (1991) and is typically constructed by taking the share of employment at the region-industry level in a period prior to the sample period (the share) and multiplying it by an industry-year level measure of employment or wage growth (the shift) and summing up over all industries at the region-year level. One of the first instances of using such an instrument for labor demand can be found in Freeman (1980).

and wage models.

Second, we use an instrumental variable to mitigate the remaining bias by using the AEWR lagged one period as an instrument for the AEWR in the current period. We demonstrate in Section 3 that the model must control for labor supply shocks over time for the exclusion restriction to be satisfied. We argue that the inclusion of the year and state fixed effects adequately controls for these shocks and that the two-stage least squares estimates identify the causal effects of interest.

Our empirical analysis uses data from the National Agricultural Workers Survey to measure wages and total hours of work per year for domestic farmworkers. As a supplement, we use data from the Quarterly Census of Employment and Wages to generate a measure of full-time-equivalent employment (FTE) of domestic workers. In a subsequent analysis, we subtract H-2A FTE employment from the QCEW employment measures for states that include H-2A employment in their QCEW data. The H-2A FTE employment data is calculated from the number of jobs certified and the length of contracts in the DOL's H-2A disclosure database. The AEWR data are obtained from the USDA's Farm Labor Survey through NASS Quickstats.

Our preliminary results indicate that a 10% increase in the real (i.e., inflation adjusted) AEWR causes a 5.4% increase in real domestic wages nationwide and a 3.3% increase in the top 10 H-2A employment states. These results are consistent with those in Buccola, Li, and Reimer (2012) and produce results that are qualitatively similar to the hourly wage results of Moretti and Perloff (2000). We also find some estimates that suggest a 10% increase in the AEWR could cause as much as a 21% increase in the number hours worked for domestic employees.

Our study makes several contributions. First, we isolate the causal effects of the AEWR on domestic farmworker labor market outcomes. In the extant literature, minimum wages are exogenously-imposed policy variables with readily estimatable impacts on labor market outcomes. In our case, however, the AEWR is endogenous to current labor-market conditions. That is, the AEWR is based on a measure of lagged wages, which is endogenous if labor market shocks in one period impact labor market outcomes in the following period. To the best of our knowledge, only a few studies have analyzed the impacts of minimum wages in the US agricultural sector, none of which have produced estimates of the causal effects of the AEWR on wages or employment (Ifft, 2021; Buccola, Li, and Reimer, 2012; Kandilov and Kandilov, 2020; Meer and West, 2016; Moretti and Perloff, 2000).

Second, we contribute to the recent farm labor literature by examining linkages between the H-2A and domestic labor markets. In doing so, we uncover an important structural relationship between the wages guaranteed to H-2A workers and wages of domestic farm employees. Moreover, we investigate whether such effects are heterogeneous among different subpopulations of the farm workforce and find that workers who are less vulnerable are likely more equipped to leverage the AEWR as benchmark in their wage and employment negotiations with employers. This results holds when focusing on workers who are documented, have good English skills, better educational attainment, male, and are not hired by farm labor contractors.

Last, we contribute to the policy discussion regarding the AEWR calculation method by providing insights into the unintended consequences of changes to the AEWR calculation method. We contribute to this discussion by providing a quantitative measure of potential externalities using an example from a recently proposed policy change that would have frozen the AEWR for a year. Our findings suggest that the AEWR likely influences domestic farmworker wages, so it is not the neutral benchmark it is intended to be. As a result, any changes made to the AEWR methodology could have a significant impact on agricultural employers and the domestic workers they employ.

The following section provides some background details related to the H-2A visa program and the AEWR. Section 2 provides a simple theoretical framework to investigate whether the AEWR effects are expected to be positive or negative on domestic labor market outcomes, Section 3 describes out empirical strategy and data, and Section 4 describes the results. We provide some concluding remarks in Section 5.

1 Background

In 1952, the H-2 program was initiated by the Immigration and Nationality Act, permitting foreign laborers to enter the country on a temporary basis to perform "low-skilled labor" in both the agricultural and non-agricultural sectors. With the passage of the Immigration Reform and Control Act in 1986, the H-2 program was broken up into the H-2A program for agricultural workers and the H-2B program for non-agricultural workers. Currently there is no cap on the number of visas that can be issued, but agricultural employers must certify that they are unable to find domestic workers before they are approved to hire H-2A workers. Foreign agricultural workers present in the country under the H-2A visa program must leave the US once their visas expire.

Over the past decade, the farm labor supply has become tighter due to a number of political, economic, and demographic factors, and the H-2A program has rapidly expanded to fill the void. Between FY2012 and FY2022, the number of H-2A jobs certified to agricultural employers increased by more than 300% from about 85,000 to over 370,000 (see Figure 1; USDOS, 2021).[6] In 2020, the DOL certified agricultural employers to fill about 10 percent of the full-time equivalent (FTE) jobs on US crop farms with H-2A guest workers, accruing an estimated H-2A wage bill of about $3.5 billion (Castillo, Martin, and Rutledge, 2022).

Low-skilled foreign-born workers tend to have low reservation wages and have been viewed as an economic threat to the domestic workforce (Congressional Research Service, 2008). In an attempt to mitigate adverse effects from the employment of temporary foreign workers in the agricultural sector, H-2A workers and the domestic workers who work for H-2A employers must be paid an amount no less than the AEWR. They must also be paid the highest of the state or federal minimum wage, the prevailing wage, or the state AEWR.

[6] Historically between 70 and 80 percent of the jobs certified by the DOL have actually been issued a visa by the DOS.

Figure 1: Number of H-2A Jobs Certified, FY2005 – FY2022

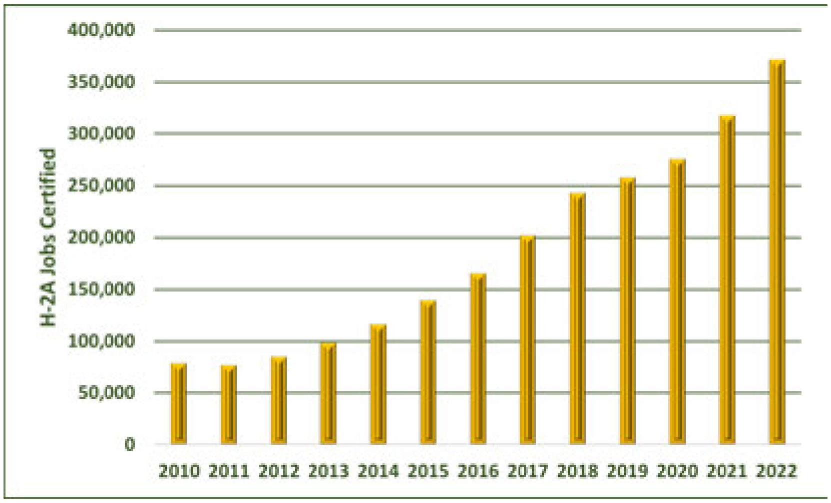
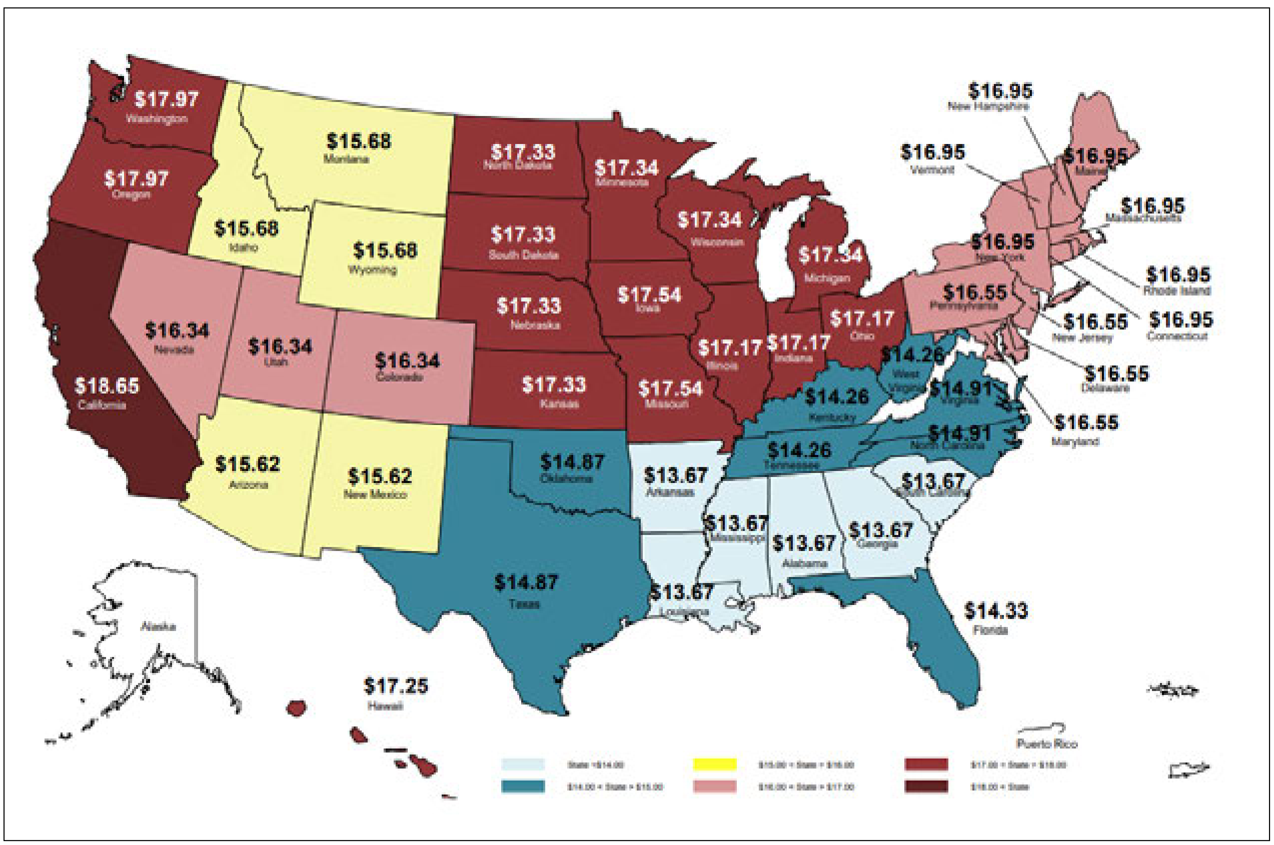


Figure 2: Adverse Effect Wage Rates in 2023



Source: https://www.dol.gov/sites/dolgov/files/ETA/oflc/pdfs/AEWR-Map-2023.pdf.

The state-level AEWRs are currently based on data from the USDA's Farm Labor Survey (FLS), which "provides the basis for employment and wage estimates for all workers directly hired by U.S. farms and ranches (excluding Alaska)" (NASS, 2021). In 2023, the AEWRs ranged from a low of \$13.67 in the southeastern part of the country to a high of \$18.65 in California (see Figure 2). The AEWRs are adjusted on an annual basis (typically upward) and are supposed to reflect the average wage in the region from the previous year.

According to DOL estimates, an AEWR freeze would save employers of H-2A workers an estimated \$140 million a year (DOL, 2020). Castillo, Martin, and Rutledge (2022) estimate that an AEWR freeze could also save employers an additional \$29 million per year for the corresponding US workers who are employed by the H-2A employers.[7] In addition to the economic impacts associated with the direct employment of H-2A workers, any proposed changes to the AEWR could potentially save farm employers hundreds of millions of dollars if they also slow the wage growth of domestic farmworkers, who make up about 85% of the farm workforce.

2 Theory

To gain insight into the theoretical underpinnings of the relationship between the AEWR and domestic farmworker employment and wages, we utilize a simple cost minimization framework. Formally, we assume that there is an aggregate farmer in a state that produces a certain amount of a crop output using two inputs, domestic labor (D) and H-2A labor (H). We assume that the farmer's objective is to minimize total labor costs subject to a given amount of production, which is generated from a Cobb-Douglas production technology.

To make ideas clear, we assume that the supply of H-2A labor H_s is perfectly elastic at the level of the AEWR, which is set by the government prior to the growing season. We define the wage level along the domestic supply (resp. demand) schedule as w_s^D (resp. w_d^D), where the capital superscript D denotes "domestic" and the lowercase subscripts s and d

[7] This figure was calculated by first identifying the number of jobs that were requested in each partially approved H-2A application that were not granted. Then the number of jobs in each contract was multiplied by the value of the contract specified in the application. Finally the total value was calculated by summing up the value over all jobs.

denote the supply and demand, respectively. We drop the superscripts on the H-2A wage variable (the AEWR) and simply define it as w^H, where the capital superscript H refers to H-2A. Further, we assume that the domestic and H-2A labor markets clear in the current growing season, such that $D_s = D_d$ and $H_s = H_d$. We define the labor supply function as follows:

$$D_s = (w_s^D)^\gamma \iff \ln D_s = \gamma \ln w_s^D, \tag{1}$$

where $\gamma \geq 0$. We characterize the farmer's labor input decision making process in the current growing as

$$\begin{aligned} &\min_{D,H} \; w^D D + w^H H \\ &\text{s.t. } \; Q \leq A D^\alpha H^\beta, \end{aligned}$$

where $\alpha > 0, \beta > 0$, and $\alpha + \beta \leq 1$. Using this framework, the Lagrangean function can be expressed as follows:

$$\mathscr{L} = w^D D + w^H H - \lambda[Q - A D^\alpha H^\beta].$$

The first order conditions imply that

$$H = \frac{\beta w^D}{\alpha w^H} D. \tag{2}$$

2.1 AEWR Effects on Domestic Farmworker Wages

Substituting (2) into the production function, taking logs, and solving for $\ln D_d$ gives us the optimal demand for domestic labor:

$$\ln D_d = \frac{1}{\alpha+\beta} \ln\left(\frac{Q}{A}\right) + \frac{\beta}{\alpha+\beta} \ln\left(\frac{\alpha}{\beta}\right) + \frac{\beta}{\alpha+\beta} \ln w^H - \frac{\beta}{\alpha+\beta} \ln w_d^D \tag{3}$$

Since the labor markets clear, we can use equations (1) and (3) to set $\ln D_d = \ln D_s$ to derive the following equilibrium relationship:

$$\gamma \ln w_s^D = \frac{1}{\alpha+\beta} \ln\left(\frac{Q}{A}\right) + \frac{\beta}{\alpha+\beta} \ln\left(\frac{\alpha}{\beta}\right) + \frac{\beta}{\alpha+\beta} \ln w^H - \frac{\beta}{\alpha+\beta} \ln w_d^D.$$

In equilibrium $w_s^D = w_d^D$, so after rearranging terms and dropping the s and d subscripts on the w^D variables, the following equation can be derived for the equilibrium domestic worker wage:

$$\ln w^{D*} = \frac{1}{\gamma\alpha+\gamma\beta+\beta} \ln\left(\frac{Q}{A}\right) + \frac{\beta}{\gamma\alpha+\gamma\beta+\beta} \ln\left(\frac{\alpha}{\beta}\right) + \frac{\beta}{\gamma\alpha+\gamma\beta+\beta} \ln w^H. \tag{4}$$

As can be seen from equation (4), for a given technology and output level, the elasticity of domestic farmworker wages with respect to the AEWR is non-negative. That is:

$$\Lambda \equiv \frac{\partial \ln w^{D*}}{\partial \ln w^H} = \frac{\beta}{\gamma\alpha+\gamma\beta+\beta} \geq 0.$$

The condition $\Lambda \geq 0$ results from the fact that, for a given production technology and output level, an increase in the AEWR induces a decrease in H-2A employment (i.e., a movement along the H-2A labor demand curve), which requires an increase in domestic employment to maintain production. Thus, the demand for domestic workers increases, resulting in an increase in domestic worker wages.

2.2 AEWR Effects on Domestic Farmworker Employment

From equations (1) and (3), we can derive the following inverse demand and inverse supply equations for domestic labor:

$$\ln w_d^D = \frac{1}{\beta} \ln\left(\frac{Q}{A}\right) + \ln\left(\frac{\alpha}{\beta}\right) + \ln w^H - \frac{\alpha+\beta}{\beta} \ln D_d \tag{5}$$

and

$$\ln w_s^D = \frac{1}{\gamma} \ln D_s. \tag{6}$$

By setting (5) equal to (6) and dropping the subscripts on the D variables, one can derive the following equation for the equilibrium employment of domestic workers:

$$\ln D^* = \frac{\gamma}{\gamma\alpha + \gamma\beta + \beta} \ln\left(\frac{Q}{A}\right) + \frac{\gamma\beta}{\gamma\alpha + \gamma\beta + \beta} \ln\left(\frac{\alpha}{\beta}\right) + \frac{\gamma\beta}{\gamma\alpha + \gamma\beta + \beta} \ln w^H.$$

Thus, for a given production technology and level of output, as shown below, the elasticity of domestic farmworker employment with respect to the AEWR is non-negative.

$$\Omega \equiv \frac{\partial \ln D^*}{\partial \ln w^H} = \frac{\gamma\beta}{\gamma\alpha + \gamma\beta + \beta} \geq 0.$$

The condition $\Omega \geq 0$ arises because, for a given technology and output, an increase in the AEWR will induce substitution out of H-2A employment into the use of domestic workers. Thus, holding constant output and technology, an increase in the AEWR should lead to higher domestic employment. In the following section, we describe our empirical model that is used to investigate whether $\Lambda \geq 0$ and $\Omega \geq 0$ and quantify the magnitudes.

3 Empirical methodology and data

3.1 Domestic Wage Model

Our primary research objective is to test the null hypothesis that the AEWR has no effect on the wages of domestic farmworkers against the alternative hypothesis that there is an effect. From an empirical perspective, the main identification challenge is overcoming the issue of omitted variables bias. To help mitigate bias that results from geographic differentials in the accumulation of human capital, our dependent variable identifies the average wage at the state-year level of aggregation after controlling for relevant human capital variables using the method of Reed and Danziger (2007).[8] Although this approach alleviates bias from individual characteristics, the fact that the AEWR is a measure of lagged wages presents additional challenges. To make ideas clear, suppose we want to estimate the following model:

$$\ln w_{st}^{D} = \Lambda \ln w_{st}^{H} + \phi_s + \phi_t + \epsilon_{st},$$

where $\ln w_{st}^{D}$ identifies the average log real wage (in \$2020) of domestic farmworkers in state s in survey year t (net of individual-level observables), $\ln w_{st}^{H}$ identifies the natural logarithm of the real AEWR (in \$2020), ϕ_s are state fixed effects, ϕ_t are year fixed effects, and ϵ_{st} is the error term. Because the AEWR is based on wage data from the previous period, it is likely influenced by labor supply and demand shocks in period $t-1$. If labor market shocks from the previous period are correlated with the labor market outcomes in the current period, then, to the extent that these shocks are not controlled for in the model, the OLS estimates would be biased.

To formalize this idea, suppose the researcher is able to adequately control for labor demand shocks in period $t-1$ such that the error term only contains an omitted labor supply shock variable from the previous period and an idiosyncratic error term such that the true model is

$$\ln w_{st}^{D} = \Lambda \ln w_{st}^{H} + \phi_t + \phi_s + \beta_1 LD_{st-1} + \underbrace{\beta_2 LS_{st-1} + \nu_{st}}_{\epsilon_{st}}, \tag{7}$$

where LD_{t-1} (respectively LS_{t-1}) denotes a labor demand (respectively supply) shock variable in period $t-1$, and ν_{st} is the error term that satisfies the condition $\mathbb{E}[\nu|LS, LD] = 0$. If labor market shocks are serially correlated with domestic wages, then it would follow that

[8] See Appendix A for details about the approach used to control for human capital accumulation.

$\beta_1 \geq 0$ and $\beta_2 \leq 0$.[9] Since the AEWR in period t is determined by the market conditions in year $t-1$, it is a function of lagged labor supply and demand shocks in period $t-1$, so it can be modeled as follows:

$$\ln w_{st}^{H} = \delta_1 LD_{st-1} + \delta_2 LS_{st-1} + \xi_{st}, \tag{8}$$

where $\mathbb{E}[\xi|LS, LD] = 0$. Because labor demand (respectively supply) shocks in period $t-1$ will tend to increase (respectively decrease) regional wages in period $t-1$ (and thus the AEWR in period t), it is natural to assume that $\delta_1 \geq 0$ and $\delta_2 \leq 0$. Under the assumption that labor demand shocks are not correlated with labor supply shocks (i.e., $\text{cov}(LS_t, LD_{t-k}) = 0$), where $k \in (0, ..., T-1)$, the probability limit of the OLS coefficient on the log AEWR variable in Equation (7) can be expressed as follows:[10]

$$\Lambda^{OLS} = \Lambda + \overbrace{\beta_2 \delta_2 \text{var}(LS_{st-1})}^{\text{Bias} \geq 0}.$$

To help mitigate the potential for upward bias resulting from omitted labor demand shocks, we develop a proxy for agricultural sector labor demand shocks using a Bartik instrument and include it as a control variable (Basso and Peri, 2015; Notowidigdo, 2020; Bartik, 1991). Our Bartik control variable is defined as follows:

$$LD_{st-1} = \sum_j \left[\left(\frac{emp_{js,1990}}{emp_{s,1990}} \right) \Delta \ln w_{jt-1} \right],$$

where j denotes one of five agricultural sectors defined by NAICS codes 111, 112, 113, 114, and 115, $emp_{js,1990}$ denotes the employment in sector j in state s in 1990, $emp_{s,1990}$ denotes total agicultural employment in the state in 1990 such that the term in parentheses represents

[9] If labor supply and demand shocks are not serially correlated with wages (i.e., $\beta_1 = \beta_2 = 0$), there would be no omitted variables bias. In that case, the OLS estimate would identify the causal effect of interest. As a result, controlling for labor demand shocks should reduce the magnitude of the positive regression coefficients.

[10] Note that the omitted variables bias from unobserved lagged labor demand shocks can be expressed as $\beta_1 \delta_1 \text{var}(LD_{st-1}) \geq 0$.

the share of sector j's ag employment in the state in 1990, and $\Delta \ln w_{jt-1}$ denotes the change in log average wage in sector j across the entire country between 1990 and year $t-1$.[11] Nevertheless, despite controlling for lagged labor demand shocks by including LD_{st-1} as a control variable, an OLS estimate of Λ would still be biased upwards due to the unobserved lagged labor supply shock variable.

To further mitigate this bias, we deploy an instrumental variable using a lagged log AEWR variable (i.e., $\ln w^H_{st-1}$). In order for this instrument to satisfy the exclusion restriction, assuming that labor supply and demand shocks are not correlated, it must be the case that $\text{cov}(\ln w^H_{st-1}, \varepsilon_{st}) = 0 \iff \text{cov}(\ln w^H_{st-1}, \beta_2 LS_{t-1}) = 0 \iff \text{cov}(\delta_1 LD_{st-2} + \delta_2 LS_{st-2}, \beta_2 LS_{st-1}) = 0 \iff \delta_2\beta_2\text{cov}(LS_{st-2}, LS_{st-1}) = 0$. In other words, either (i) labor supply shocks from the previous period do not have an impact on wages in the current period (i.e., $\beta_2 = 0$ and there are no omitted variables), (ii) there is no serial correlation between labor supply shocks in a state, or (iii) serial correlation between labor supply shocks is adequately controlled for in the model. We argue that (iii) is plausible because (a) our model includes year fixed effects, which control for labor supply shocks that are common to all states within a given year such that macroeconomic labor supply shocks that impact regional labor markets across time are effectively controlled for and (b) to the extent that some states are more susceptible to labor supply shocks, such that a state that experiences a shock in one year would tend to experience a similar shock in the following year, the state fixed effects adequately control for that. Thus, it is plausible that, conditional on our set of year and state fixed effects and the Bartik labor demand control variable, our preferred two-stage least squares estimate identifies the causal effect of interest.

[11] We demonstrate in Section 4 that the inclusion of the Bartik control variable does, in fact, mitigate upward bias in both the wage and employment models.

3.2 Domestic Employment Model

Suppose we want to estimate the following model:

$$\ln D_{st} = \Omega \ln w_{st}^{H} + \phi_s + \phi_t + \psi_{st},$$

where $\ln D_{st}$ identifies the log employment of domestic farmworkers in state s in year t, $\ln w_{st}^{H}$ identifies the natural logarithm of the real AEWR (in \$2020), ϕ_s are state fixed effects, ϕ_t are year fixed effects, and ψ_{st} is the error term. Following the rationale from Section 3.1, suppose we are able to adequately control for labor demand shocks using a Bartik instrument such that the error term contains unobserved labor supply shocks from period $t-1$ and the true model is

$$\ln D_{st} = \Omega \ln w_{st}^{H} + \phi_t + \phi_s + +\kappa_1 LD_{st-1} + \underbrace{\kappa_2 LS_{st-1} + \varepsilon_{st}}_{\psi_{st}}, \tag{9}$$

where LD_{t-1} and LS_{t-1} are defined in Section 3.1. Because labor supply and demand shocks are both positively correlated with employment, to the extent that labor supply shocks in the previous period influence domestic employment in the current period, it is plausible that $\kappa_1 \geq 0$ and $\kappa_2 \geq 0$.[12] Recall that the AEWR is modeled as a function of lagged labor supply and demand shocks in equation (8) such that $\delta_1 \geq 0$ and $\delta_2 \leq 0$. If labor demand shocks are not systematically correlated with labor supply shocks (i.e., $\text{cov}(LS_t, LD_{t-k}) = 0$, where $k \in \{0, ..., T-1\}$), the probability limit of the OLS coefficient on the log AEWR variable in Equation (9) can be expressed as follows:

$$\Omega^{OLS} = \Omega + \overbrace{\kappa_1 \delta_1 \text{var}(LS_{st-1})}^{\text{Bias } \leq 0}.$$

[12]If labor supply shocks in the previous period do not impact wages in the current period, then $\kappa_2 = 0$, and the OLS estimate would identify the causal effect.

As a result, if labor demand shocks are adequately controlled for, the sign of the omitted variables bias in the employment model is negative.[13] Recall from section 3.1 that we use a Bartik instrument as a proxy for labor demand and find empirical results that are consistent with theoretical expectations. The condition for the exclusion restriction to hold in the employment specification is the same as the wage specification, so our argument for identification remains unchanged.

3.3 Data

We bring together data from five sources to conduct our analyses. Our domestic farmworker wage data consists of individual-level data from the 1990-2020 samples of the National Agricultural Workers Survey (NAWS). We restrict our sample to include only those individuals who were between the ages of 18 and 64 at the time of the survey, which retains about 95% of the sample. We adjust the nominal wage data to real 2020 dollar values using a consumer price index. The NAWS also contains a host of individual-level variables that we use as controls to generate a "residualized" wage variable, which is used as our outcome variable in the wage analysis (see Appendix A for more details).

Our Bartik control variable is constructed from average annual employment and weekly wage data from the Quarterly Census of Employment and Wages (QCEW). To create a proxy for labor demand shocks in the agricultural sector, we utilize data for each of the three-digit industries within the agricultural sector. Specifically, we use NAICS codes 111 (crop production), 112 (animal production), 113 (forestry and logging), 114 (fishing, hunting, and trapping), and 115 (crop support services). The employment data is aggregated at the state-industry level using 1990 as the base year, and the wage data is aggregated at the industry-year level.

We use two measures of domestic employment. The first measure identifies extensive

[13] Note that the omitted variables bias from an unobserved labor demand shock variable would be $\kappa_1\delta_1\text{var}(LD_{st-1}) \geq 0$. Therefore, if labor demand shocks are not adequately controlled for, the overall sign of the bias would be ambiguous.

margin employment in terms of the number of full-time equivalent (FTE) jobs from the QCEW. This measure of employment is identified by the sum of annual average employment from the direct hire crop (NAICS 111) and crop support (NAICS 1151) sectors. As a robustness test for extensive margin employment, we subtract the number of certified FTE H-2A jobs from the QCEW measures for California, Oregon, and Washington and re-run the analysis.[14] We use H-2A job certification data from the DOL's disclosure data. As a robustness check, we subtract the number of H-2A FTEs certified from the QCEW FTE measures for the states of California, Oregon, and Washington, who include H-2A workers in their data. For a given certified H-2A contract in a state, the number of H-2A FTEs is calculated by determining the length of the contract, dividing by 365, and multiplying by the number of workers certified in the contract. Then we sum up over all the contracts in a state.

The second measure identifies intensive margin employment by using the average number of hours worked per year per employee from the NAWS. The number of hours worked is calculated by taking the product of the number of hours of farm work performed during the previous week and the total number of weeks of farm work performed during the previous 52 weeks for each worker.

Finally, the AEWR data were obtained from the USDA's Farm Labor Survey (FLS) through the NASS Quickstats website. Specifically, the AEWR represents the average regional wage for hired crop and animal workers from the previous year, which we assign to the states that belong to each region. A selection of summary statistics can be found in Table 1.

[14]The H-2A data is only available as far back as 2008, so this procedure significantly reduces the sample size, so we urge caution when interpreting these results.

Table 1: Summary Statistics

	Mean	St. Err. of Mean	Observations
Real AEWR ($2020)	11.451	1.274	684
Real wage ($2020)	10.961	0.026	61,432
Age (years)	34.893	0.095	61,432
Male	0.752	0.004	61,432
Married	0.586	0.004	61,432
Undocumented	0.450	0.004	61,432
Speaks good English	0.251	0.004	61,432
Number of years of education	7.641	0.032	61,432

4 Results

4.1 Domestic Farmworker Wage Analysis

4.1.1 Main Analysis

Our main results from the wage analysis are presented in Table 2. The odd numbered columns contain the coefficients from the OLS regressions, and the even numbered columns contain the estimates from the 2SLS regressions using the lagged AEWR variable as an instrument. The results from our preferred specification are displayed in column (8), which include the state and year fixed effects, as well as the Bartik control variable. In cases where the IV coefficients are not statistically significant and the OLS coefficients are, due to the omitted variables bias from the lagged labor supply shocks, the OLS coefficient can be interpreted as an upper bound for the effect of interest. While OLS bounds for the population parameters could provide useful information, they are less informative than a well-defined point estimate, so we do not rely upon them for inference.

The top panel in Table 2 shows the results for the entire sample of states, while the bottom panel shows the results from a sample that includes only the leading top 10 H-2A employment states.[15] First, it is worth noting that the inclusion of the Bartik control variable reduces the magnitudes of the empirical estimates, suggesting that controlling for

[15] The top 10 H-2A employment states include California, Florida, Washington, Oregon, Texas, North Carolina, Louisiana, Michigan, Arizona, and Georgia.

Table 2: Estimates of the Effects of the AEWR on Domestic Farmworker Wages

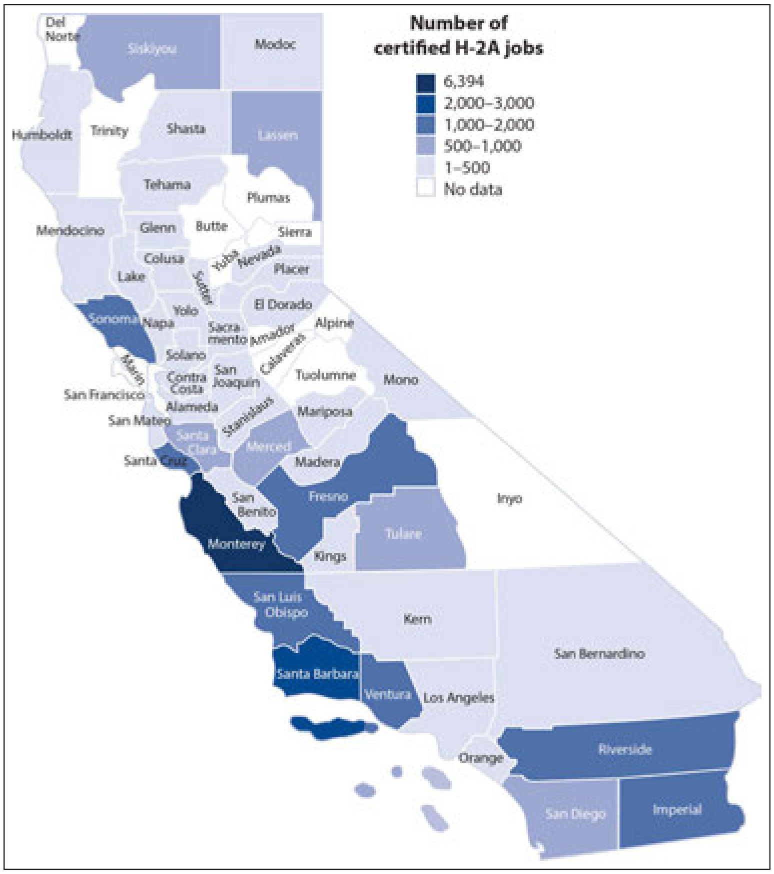
	(1) OLS $\ln w^D$	(2) IV $\ln w^D$	(3) OLS $\ln w^D$	(4) IV $\ln w^D$	(5) OLS $\ln w^D$	(6) IV $\ln w^D$	(7) OLS $\ln w^D$	(8) IV $\ln w^D$
	All States							
$\ln w^H$	0.815***	0.810***	0.642***	0.620***	0.568***	0.380	0.586***	0.536**
	(0.061)	(0.064)	(0.116)	(0.124)	(0.145)	(0.239)	(0.155)	(0.250)
N	684	684	684	684	684	684	684	684
	Top 10 H-2A States							
$\ln w^H$	0.891***	0.880***	0.787***	0.759***	0.584***	0.372*	0.430**	0.325*
	(0.091)	(0.092)	(0.160)	(0.168)	(0.159)	(0.207)	(0.171)	(0.167)
N	276	276	276	276	276	276	276	276
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.
* $p < .1$, ** $p < .05$, *** $p < .01$

labor demand shocks does, in fact, mitigate upward bias. For example, a comparison of the IV coefficient in column (2) to the one in column (4) of the top panel reveals a reduction in the elasticity estimate from 0.81 to 0.62. We also note that in all cases, the IV estimates are smaller than the OLS estimates, suggesting that our instrument helps resolve upward bias from the unobserved labor supply shocks. Comparing the coefficient in column (8) (our preferred specification) to that of column (4) indicates that additional upward bias from unobserved labor supply shocks is mitigated once unobserved macroeconomic shocks in a given year and state-specific factors are controlled for.

Our preferred specification in column (8) reveals an elasticity of domestic farmworker wages with respect to the AEWR of 0.54. When focusing on the top 10 H-2A employment states, our estimate is 0.33, which is smaller than the estimate from the entire sample. Such a result could arise from the fact that the top H-2A states also have a significantly larger number of domestic employees such that the AEWR is relevant benchmark for a relatively smaller share of the domestic workforce, possibly in the few areas where H-2A employment is more prevalent. For example, in FY2020, H-2A employment in California only comprised

Figure 3: Number of H-2A Jobs Certified in California Counties During FY2020



Source: https://calag.ucanr.edu/archive/?article=ca.2021a0020.

three percent of full time equivalent agricultural employment. Moreover, the majority of H-2A employment in California was concentrated in just a few counties, and the H-2A share of total agricultural employment in those counties was relatively small (see Figure 3).

4.1.2 Heterogeneous Effects

Next, we turn our attention to various sub-populations of the domestic farm workforce to investigate whether there is evidence of heterogeneity in the AEWR effects between different groups of workers. We select a set of key observable farmworker characteristics that are arguably associated with more and less vulnerability. For example, a number of studies have found evidence that labor markets are segmented with respect to documented and undocumented status and that undocumented workers tend to experience significant wage disparities even after controlling for differences in human capital accumulation (Borjas and Cassidy, 2019; Durand, Massey, and Pren, 2016; Massey and Gentsch, 2014; Rivera-Batiz, 1999, e.g.,). These workers are often assigned to job tasks that have little to no potential for upward mobility (Taylor, 1992). As such, it is of interest to investigate whether undocumented workers are able to equally leverage the AEWR as a benchmark when attempting to negotiate higher wages. Evidence to the contrary would suggest that they have a weaker bargaining position relative to that of documented workers, perhaps due to the relatively smaller set of available employment options.

Table 3: Estimates of the Effects of the AEWR on Domestic Farmworkers Wages by Legal Status and English Ability

	(1) OLS $\ln w^D$	(2) IV $\ln w^D$	(3) OLS $\ln w^D$	(4) IV $\ln w^D$	(5) OLS $\ln w^D$	(6) IV $\ln w^D$	(7) OLS $\ln w^D$	(8) IV $\ln w^D$
	Documented							
$\ln w^H$	0.830***	0.830***	0.703***	0.696***	0.689***	0.638***	0.660***	0.558*
	(0.060)	(0.064)	(0.115)	(0.125)	(0.147)	(0.228)	(0.183)	(0.285)
N	674	674	674	674	674	674	674	674
	Undocumented							
$\ln w^H$	0.883***	0.872***	0.603***	0.563***	0.492***	0.125	0.393**	0.394*
	(0.088)	(0.089)	(0.131)	(0.142)	(0.175)	(0.303)	(0.162)	(0.217)
N	513	513	513	513	513	513	513	513
	Speaks Good English							
$\ln w^H$	0.809***	0.819***	0.762***	0.779***	0.630***	0.708***	0.626***	0.612**
	(0.062)	(0.063)	(0.119)	(0.126)	(0.167)	(0.215)	(0.212)	(0.300)
N	642	642	642	642	642	642	642	642
	Does Not Speak Good English							
$\ln w^H$	0.777***	0.764***	0.535***	0.493***	0.493***	0.154	0.396***	0.395*
	(0.089)	(0.090)	(0.137)	(0.146)	(0.156)	(0.231)	(0.144)	(0.230)
N	560	560	560	560	560	560	560	560
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.

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

Table 3 displays the result for documented and undocumented workers in the top and

Table 4: Estimates of the Effects of the AEWR on Wages of Domestic Farmworkers by by Employer Type, Education Level, and Gender

	(1) OLS $\ln w^D$	(2) IV $\ln w^D$	(3) OLS $\ln w^D$	(4) IV $\ln w^D$	(5) OLS $\ln w^D$	(6) IV $\ln w^D$	(7) OLS $\ln w^D$	(8) IV $\ln w^D$
	Hired Directly by a Farmer							
$\ln w^H$	0.816***	0.814***	0.646***	0.628***	0.574***	0.402*	0.623***	0.623***
	(0.056)	(0.059)	(0.114)	(0.121)	(0.138)	(0.212)	(0.139)	(0.203)
N	682	682	682	682	682	682	682	682
	Hired by a Farm Labor Contractor							
$\ln w^H$	0.958***	0.933***	0.671***	0.592***	0.500	-0.079	0.227	-0.652
	(0.142)	(0.133)	(0.163)	(0.159)	(0.366)	(0.666)	(0.467)	(1.202)
N	169	169	169	169	169	169	169	169
	High School Education							
$\ln w^H$	0.783***	0.791***	0.643***	0.650***	0.587***	0.615*	0.572**	0.521
	(0.062)	(0.066)	(0.125)	(0.133)	(0.205)	(0.325)	(0.226)	(0.343)
N	653	653	653	653	653	653	653	653
	Less than High School Education							
$\ln w^H$	0.779***	0.778***	0.623***	0.610***	0.387**	0.115	0.236	0.134
	(0.077)	(0.075)	(0.121)	(0.126)	(0.191)	(0.258)	(0.195)	(0.278)
N	625	625	625	625	625	625	625	625
	Male							
$\ln w^H$	0.831***	0.825***	0.659***	0.634***	0.498***	0.228	0.506**	0.365
	(0.072)	(0.076)	(0.126)	(0.134)	(0.180)	(0.353)	(0.191)	(0.412)
N	676	676	676	676	676	676	676	676
	Female							
$\ln w^H$	0.677***	0.668***	0.594***	0.568***	0.371**	0.129	0.481***	0.304
	(0.068)	(0.070)	(0.138)	(0.154)	(0.164)	(0.325)	(0.145)	(0.306)
N	534	534	534	534	534	534	534	534
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.
* $p < .1$, ** $p < .05$, *** $p < .01$

second panels and for those workers who speak good English and those who do not in the bottom two panels. Our preferred estimates in column (8) reveal that the AEWR has a larger effect on the wages of documented workers. The elasticity for documented workers is 0.56 while it is only 0.39 for undocumented workers. This evidence is consistent with a scenario where less vulnerable workers are more capable of leveraging the threat of seeking work outside their current place of employment unless they receive a wage that is comparable to the AEWR. With respect to workers who speak good English, the results are qualitatively similar. The elasticity for workers who speak good English is 0.61 while it is only 0.40 for those who do not.

The results in Table 4 compare the effects for workers who are hired directly by farmers to those who are brought to farms by farm labor contractors in the top two panels. In the third and fourth panels, we compare the effects on workers who do and do not have a a high school education (i.e., 12 years of education or more). Interestingly, the elasticity for workers directly hired by producers is about 0.62 while the estimate for those working for farm labor contractors is statistically insignificant. With respect to education, our preferred estimates are both statistically insignificant, but the point estimates for workers with a high school education are consistently larger across the entire set of results, suggesting that the effects are more pronounced for better educated workers. The OLS estimate in column (7) suggests that the elasticity could be as large as 0.57 for domestic employees with a high school education. Finally, we compare male and female workers. While the IV coefficients are not significant in our preferred specification, again, the coefficients across the entire set of results reveal that the effect is likely larger for male workers, consistent with a weaker bargaining position among female workers.

4.2 Domestic Farmworker Employment Analysis

4.2.1 Extensive Margin Employment

Next, we turn our attention to the employment effects of the AEWR. In this section, we focus on the impacts on extensive margin employment measured by full-time equivalent jobs. Table 5 displays the results for the full 1990-2020 sample. While all the coefficients are positive, which is consistent with our theoretical expectations, this set of results are generally not statistically significant, suggesting that changes in the AEWR do not have a robust impact on the number of jobs being filled by domestic employees. This result holds when we focus on workers in all states, as well as those working in the top 10 H-2A employment states.

Table 6 displays the results from a restricted sample period during which we have data on the number of FTE H-2A jobs. Some states include H-2A workers in their QCEW employment measures, so the results in Table 5 may not accurately reflect domestic employment for those states. We know this is true for California, Oregon, and Washington, and that it is not true for Florida. Therefore, we subtract the number of FTE H-2A jobs certified from the QCEW FTE measures in these states to obtain a more accurate measure of domestic employment. While the estimate for the entire US is insignificant, our preferred estimate for the top 10 H-2A states is statistically significant, suggesting that a 10% percent increase in the AEWR causes a similar increase in the number of domestic jobs in those states, holding constant other factors. However, due to the small sample size and instability of coefficients among the different model specifications in this panel, we are hesitant to interpret this estimate as causal and urge readers to view this result with caution.

Table 5: Estimates of the Effects of the AEWR on Domestic FTE Employment

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	$\ln D$	$\ln D$	$\ln D$	$\ln D$	$\ln D$	$\ln D$	$\ln D$	$\ln D$
	All States							
$\ln w^H$	2.469**	2.693**	2.250	2.656*	0.601	0.891	0.761	1.125
	(1.023)	(1.088)	(1.455)	(1.601)	(0.442)	(0.701)	(0.531)	(0.787)
N	677	677	677	677	677	677	677	677
	Top 10 H-2A States							
$\ln w^H$	3.796**	4.350***	4.179*	5.097**	0.668	0.903	0.580	0.613
	(1.571)	(1.641)	(2.095)	(2.208)	(0.626)	(0.854)	(0.726)	(0.963)
N	279	279	279	279	279	279	279	279
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.

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

Table 6: Estimates of the Effects of the AEWR on Domestic FTE Employment

	(1) OLS $\ln D$	(2) IV $\ln D$	(3) OLS $\ln D$	(4) IV $\ln D$	(5) OLS $\ln D$	(6) IV $\ln D$	(7) OLS $\ln D$	(8) IV $\ln D$
				All States				
$\ln w^H$	2.819**	2.849**	2.351	2.401	0.088	-0.210	0.330	0.124
	(1.251)	(1.320)	(1.470)	(1.573)	(0.188)	(0.769)	(0.274)	(0.787)
N	331	331	331	331	331	331	331	331
				Top 10 H-2A States				
$\ln w^H$	4.058**	4.511**	3.849	4.472**	0.228	0.178	0.697*	0.967**
	(1.754)	(1.796)	(2.195)	(2.271)	(0.342)	(0.602)	(0.335)	(0.473)
N	126	126	126	126	126	126	126	126
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.
* $p < .1$, ** $p < .05$, *** $p < .01$

4.2.2 Intensive Margin Employment

In this section, we focus our attention on intensive margin employment in the domestic workforce. Specifically, the outcome variable identifies the average number of hours worked per year by domestic workers in a state after controlling for potential selection bias from differences in human capital accumulation (see Appendix A). The results in Table 7 indicate that the elasticity of hours worked per domestic employee with respect to the AEWR is statistically significant and positive when focusing on the entire US. Our preferred estimate indicates that the elasticity is about 2.1, suggesting that higher AEWRs cause agricultural employers to stretch their domestic workforce by employing them for more hours of work throughout the year. The estimate for the top 10 counties is of a similar magnitude but is not statistically significant, possibly due to reduced estimation power from the smaller sample size.

In Appendix B we provide the intensive margin employment results for the same subsamples of workers that we investigated in Section 4.1.2. Interestingly, we find a qualitatively

Table 7: Estimates of the Effects of the AEWR on Average Hours of Work of Domestic Farmworkers

	(1) OLS $\ln D$	(2) IV $\ln D$	(3) OLS $\ln D$	(4) IV $\ln D$	(5) OLS $\ln D$	(6) IV $\ln D$	(7) OLS $\ln D$	(8) IV $\ln D$
	All States							
$\ln w^H$	1.130***	1.228***	-0.184	-0.065	1.184*	2.885***	1.670**	2.069**
	(0.300)	(0.309)	(0.386)	(0.421)	(0.643)	(1.100)	(0.624)	(0.978)
N	687	687	687	687	687	687	687	687
	Top 10 H-2A States							
$\ln w^H$	1.331**	1.495***	0.238	0.512	0.745	2.314	1.843	2.526
	(0.485)	(0.483)	(0.552)	(0.550)	(1.135)	(1.495)	(1.161)	(1.541)
N	278	278	278	278	278	278	278	278
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.
* $p < .1$, ** $p < .05$, *** $p < .01$

similar set of results. In each case, workers who tend to have characteristics that are associated with less vulnerability are influenced more by the AEWR. Specifically, workers who tend to be less vulnerable increase their duration of employment more than other workers when the AEWR increases. While the mechanism driving this result seems counter-intuitive, it is possible that this relates to the bargaining position of these workers who are able to generate higher earnings over the course of a year by securing more hours of work.

5 Conclusion

Although the use of the H-2A visa program has increased dramatically over the past two decades, the US farm labor force is largely comprised of Mexican workers who have settled in the US. For all intents and purposes, these workers are American, although many of them lack the legal authorization to reside or work in the US. Many farmworkers are weakly positioned due to their legal status and relatively low educational attainment, so they tend to be more vulnerable than other low-skilled workers.

Our simple theoretical model suggests that, for a given production technology and level of output, an increase in the wages of H-2A workers will increase the demand for domestic labor and cause domestic wages and employment to rise. Our reduced-form empirical analysis provides evidence that is consistent with this theory, substantiating the notion that changes in the AEWR have a direct effect on the labor market outcomes of domestic farmworkers. The preliminary estimate from our preferred instrumental variable model indicates that a 10% increase in the AEWR causes the average wage of domestic farmworkers to increase by 5.4% nationwide and by 3.3% in the top 10 H-2A employment states. In terms of employment, we do not find strong evidence of extensive margin adjustments, but we uncover evidence of intensive margin adjustments, suggesting that higher AEWRs induce the employment of domestic workers for longer periods of time during the course of a year. Our preliminary estimate indicates that a 10% increase in the AEWR causes a 21% increase in the average number of hours worked per domestic employee. A closer look at subsamples of the data indicate that workers who are less vulnerable tend to be more influenced by the AEWR even after controlling for various measures of human capital accumulation. As a result, it appears that workers who experience less labor market frictions are more capable of leveraging the AEWR as a benchmark in their wage and employment negotiations.

Nationally, the AEWR has grown at a rate of 3.5% per year over the past decade. Our results suggest that an AEWR freeze could potentially slow domestic farmworker wage growth by about 1.9% (3.5% $\times$ 0.54 $\approx$ 1.9%). Because domestic farm wages account for roughly \$30 billion per year, an AEWR freeze would reduce the growth of domestic worker wages by \$570 million (\$30bil $\times$ 1.9%) per year. This reduction in wage growth would be added to the estimated \$140 million in reduced wage growth for H-2A workers and the \$29 million for corresponding domestic workers who work for H-2A employers (Castillo, Martin, and Rutledge, 2022). Our findings reveal that any changes to the AEWR methodology could significantly impact the amount of compensation that domestic employees receive.

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Online appendices

A Residualized wage variables

Human capital accumulation is highly correlated with an individual's wage rate (see Jacob, 1958), so regional wage differentials may reflect differences in the composition of the population. As a result, empirical models that fail to control for individual-level characteristics tend to be biased. To alleviate concern about regional selection bias, the dependent variable used in our wage analysis is constructed by estimating a regression of the individual-level domestic farmworker wages on a full set of state-by-year fixed effects and a relevant set of individual-level observables and using the state-by-year fixed effects as the dependent variable in the final regression.[16] Our individual-level regression is defined as follows:

$$\begin{aligned}\ln O_{ist} &= y_{st} + \gamma_1 \text{Age}_{ist} + \gamma_2 \text{School}_{ist} + \gamma_3 \text{Male}_{ist} + \gamma_4 \text{Undocumented}_{ist} \\ &\quad + \gamma_5 \text{Married}_{ist} + \gamma_6 \text{English}_{ist} + \psi_{ist}, \end{aligned} \tag{A.1}$$

where $\ln O_{ist} \in \{\ln w^D_{ist}, \ln D_{ist}\}$ identifies the natural log of the real wage (in \$2020) or annual hours of employment of domestic farmworker i in state s in survey year t, $y_{st} \equiv \{\ln w^D_{st}, \ln D_{st}\}$ is a set of state-by-year fixed effects that capture the average domestic worker wages and employment after controlling for potentially confounding demographics. Using separate wage and employment outcome variables, the fixed effects from equation (A.1) are used as the dependent variables in the empirical analysis. The variable Age_{ist} is the age of the individual, School_{ist} identifies the number of years of schooling that the individual has completed, Male_{ist} is a dummy variable for being male, $\text{Undocumented}_{ist}$ is a dummy variable identifying undocumented workers, Married_{ist} is a dummy for being married, English_{ist} is a

[16] This approach follows Reed and Danziger (2007) in that we estimate region-by-year fixed effects to identify the average wage in each state in each year after controlling for potentially confounding human capital variables. Note that our approach differs slightly from Reed and Danziger (2007) in that our geographic unit of interest is the state (rather than the Metropolitan Statistical Area) and we include a more robust set of individual-level controls.

dummy variable for speaking good English, and ψ_{ist} is the error term.[17,18] We consolidate the coefficients on the variables y_{st} into a column vector that contains a single value for each state-year observation, which is used as the dependent variable in our wage analysis.

[17]Note that this is not a panel regression model. We do not observe individuals over time in the NAWS data. As a result, this regression is conducted on a sample of pooled cross sections spanning the time period 1991 to 2018.

[18]The individual-level regression is survey-design corrected according to the method prescribed by staff at the US Department of Labor through personal communication. The regression is estimated in Stata by using the "svy: regress" command after designating the "cluster" variable as the primary sampling unit, the "region12" and "cycle" variables as the two-level strata, and the variable "pwtycrd" as the probability weighting variable.

B Heterogeneous Employment Effects

Table 8: Estimates of the Effects of the AEWR on Average Hours of Work of Domestic Farmworkers by Legal Status and English Ability

	(1) OLS $\ln D$	(2) IV $\ln D$	(3) OLS $\ln D$	(4) IV $\ln D$	(5) OLS $\ln D$	(6) IV $\ln D$	(7) OLS $\ln D$	(8) IV $\ln D$
					Documented			
$\ln w^H$	0.666**	0.715***	-0.320	-0.283	1.383**	2.785***	1.713**	1.801**
	(0.266)	(0.266)	(0.380)	(0.397)	(0.607)	(0.803)	(0.650)	(0.816)
N	676	676	676	676	676	676	676	676
					Undocumented			
$\ln w^H$	2.311***	2.525***	-0.054	0.253	0.869	3.428**	1.303	2.236
	(0.421)	(0.440)	(0.529)	(0.581)	(0.804)	(1.534)	(0.912)	(1.633)
N	514	514	514	514	514	514	514	514
					Speaks Good English			
$\ln w^H$	0.388	0.433*	-0.388	-0.342	1.654**	3.390***	1.766**	2.452**
	(0.254)	(0.256)	(0.423)	(0.434)	(0.716)	(0.902)	(0.809)	(1.011)
N	646	646	646	646	646	646	646	646
					Does Not Speak Good English			
$\ln w^H$	1.916***	2.116***	0.056	0.373	0.346	2.293*	1.035	1.637
	(0.290)	(0.308)	(0.429)	(0.485)	(0.626)	(1.245)	(0.663)	(1.249)
N	559	559	559	559	559	559	559	559
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.

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

Table 9: Estimates of the Effects of the AEWR on Average Hours of Work of Domestic Farmworkers by Employer Type, Education Level, and Gender

	(1) OLS $\ln D$	(2) IV $\ln D$	(3) OLS $\ln D$	(4) IV $\ln D$	(5) OLS $\ln D$	(6) IV $\ln D$	(7) OLS $\ln D$	(8) IV $\ln D$
	Hired Directly by a Farmer							
$\ln w^H$	0.961***	1.057***	-0.328	-0.215	1.239**	3.009***	1.760***	2.377***
	(0.278)	(0.282)	(0.356)	(0.385)	(0.578)	(0.957)	(0.525)	(0.849)
N	684	684	684	684	684	684	684	684
	Hired by a Farm Labor Contractor							
$\ln w^H$	2.704***	2.703***	0.928	0.808	-0.234	-2.899	0.802	-4.896
	(0.710)	(0.789)	(1.179)	(1.380)	(1.714)	(2.087)	(2.954)	(4.162)
N	167	167	167	167	167	167	167	167
	High School Education							
$\ln w^H$	0.660**	0.757**	-0.423	-0.290	1.135	3.166**	1.104	2.109*
	(0.303)	(0.313)	(0.408)	(0.442)	(0.816)	(1.309)	(0.685)	(1.153)
N	658	658	658	658	658	658	658	658
	Less than High School Education							
$\ln w^H$	1.465***	1.561***	-0.300	-0.227	0.962	2.300**	1.070	1.260
	(0.324)	(0.338)	(0.414)	(0.491)	(0.667)	(1.168)	(0.756)	(1.212)
N	623	623	623	623	623	623	623	623
	Male							
$\ln w^H$	1.260***	1.347***	-0.142	-0.051	1.201*	2.713**	1.911***	1.998**
	(0.303)	(0.316)	(0.401)	(0.443)	(0.609)	(1.110)	(0.631)	(0.985)
N	678	678	678	678	678	678	678	678
	Female							
$\ln w^H$	1.114***	1.227***	-0.610	-0.437	0.124	1.764	-0.383	-0.231
	(0.272)	(0.302)	(0.523)	(0.584)	(0.917)	(1.226)	(1.414)	(1.940)
N	536	536	536	536	536	536	536	536
Demographic Controls	X	X	X	X	X	X	X	X
Bartik Control	–	–	X	X	X	X	X	X
State F. E.	–	–	–	–	X	X	X	X
Year F. E.	–	–	–	–	–	–	X	X

Standard errors clustered at the state level in parentheses.

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