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Do Regulations to Protect Endangered Species on Private Lands Affect Local Employment? Evidence from the Listing of the Lesser Prairie Chicken

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The U.S. Endangered Species Act is often criticized as "pitting people against wildlife" by conserving habitat at the cost of jobs, but relatively little is known about the labor market effects of listing a species under the Endangered Species Act. We examine changes in employment associated with the lesser prairie chicken, which was listed as threatened in May 2014. Using county-level employment data and variation in suitable prairie chicken habitat, we apply a difference-in-differences strategy to measure the employment effects of the listing decision. We find evidence that employment declined about 1.5% in affected counties. The effect is proportional to habitat, which means counties with relatively more habitat experienced a larger share of employment losses.

Key words: conservation, Endangered Species Act, growth, habitat

Introduction

Endangered species conservation has a controversial yet poorly understood connection to the broader economy. Species extinction rates have risen and current estimates classify one-fifth of all species as endangered, meaning those species are likely to become extinct in the near future. Without conservation, this number would be substantially higher (Hoffmann et al., 2010). Habitat modification from human activity is the greatest contributor to the decline of most species (Millennium Ecosystem Assessment, 2005). As a result, conservation policies focus on protecting endangered species habitat by i) managing public lands to serve as wildlife habitat and ii) regulating private lands through a combination of land use restrictions and incentives. Both of these policies invite controversy. In particular, land use restrictions are controversial because the costs fall disproportionately on private landowners and developers (Innes, Polasky, and Tschirhart, 1998). There is widespread public concern that protecting wildlife damages local industry and labor markets (Burke, 2004; Goodstein, 1999). Considering the scope of the conservation issue and the amount of public backlash, there is remarkably little published research quantifying the effects of endangered species regulations on local markets.

This paper contributes empirical evidence to this controversy by estimating the local employment consequences of listing an endangered species in the United States. Under the Endangered Species

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Act (ESA), species listed as endangered or threatened cannot be harmed, which includes acts that kill, injure, or significantly modify habitat essential to the species. The threat of regulatory restrictions and substantial civil and criminal penalties places a burden on landowners and industries that rely on natural resources. Many Americans fear that listing a species restricts development and reduces employment in areas with protected habitat (Goodstein, 1999; Greenstone and Gayer, 2009). We test whether this hypothesis holds for the lesser prairie chicken, whose habitat in the Great Plains intermixes with farms, ranches, and energy development. We hypothesize that employment in areas occupied by the lesser prairie chicken declined following its listing.

A large and growing research effort is investigating the economic impacts of environmental policies and environmental change using quasi-experimental methods (Greenstone and Gayer, 2009; Kuminoff, Parmeter, and Pope, 2010). Concerns about omitted variable bias have pushed empirical researchers to adopt techniques such as instrumental variables and difference-in-differences—which have a long history in public and labor economics—to identify causal relationships in economic activity (Angrist and Pischke, 2008). Recent applications in environmental economics have used these methods to identify the effects of acid rain regulations on the behavior of polluting firms (List et al., 2003; Hanna, 2010; Di Maria, Lange, and Van der Werf, 2014; Ferris, Shadbegian, and Wolverton, 2014); carbon emission regulations on low-carbon technology development (Calel and Dechezlepretre, 2016), natural amenities, and landscape change on residential property values (Horsch and Lewis, 2009; Heintzelman, 2010; Currie et al., 2015; Locke and Blomquist, 2016; Sunak and Madlener, 2016); shale gas extraction on local employment and wages (Cosgrove et al., 2015; Komarek, 2016); and farmland subsidies on the adoption of green practices and ecosystem services (Roberts and Bucholtz, 2005; Chabé-Ferret and Subervie, 2013). Recently, several papers have applied quasi-experimental methods to measure the impacts of endangered species protection on land use and development (Boškoviç and Nøstbakken, 2017; Melstrom, 2017; Wietelman and Melstrom, 2017). Our study contributes to this literature by applying difference-in-differences to measure the local labor market effects of ESA regulations.

Research on the economic effects of ESA regulations suggests that a substantial trade-off exists between species conservation and jobs. Many of these papers are unpublished and describe input–output or computable general equilibrium models to predict *ex ante* production and employment effects of impending listings (Baier and Segal, 2014; Schamberger et al., 1992) or designating critical habitat (Waters, Holland, and Weber, 1994; Watts et al., 2001). Ferris (2009) and Eichman et al. (2010) are notable departures in that they conduct econometric investigations using real-world data. Both examine changes in local employment growth following the creation of the Northwest Forest Plan to protect northern spotted owl habitat.[1] The northern spotted owl incited a national debate about the economic impacts of ESA regulations when the species was listed as threatened in 1990 (Meyer, 1997). Ferris and Eichman et al. both find evidence that the regulations restricting harvests on public land reduced local employment in the U.S. Northwest by tens of thousands of jobs. The northern spotted owl is one example of how controversial and costly endangered species protections on public lands can be.

There is a broader literature on the economics of the ESA, particularly related to preemptive habitat destruction, payments for habitat conservation, and the political economy of listing species (Langpap, Kerkvliet, and Shogren, 2018). Preemptive habitat destruction occurs when landowners manage their land in a way to keep a species off their property, in an effort to avoid land use restrictions (Innes, 2000; Shogren and Tschirhart, 2008; Zabel and Paterson, 2006). Economists have responded to this problem by designing conservation incentive programs for landowners (Langpap and Wu, 2004; Langpap, 2004, 2006; Drechsler et al., 2010) and studying the effectiveness of in situ conservation programs such as habitat conservation plans (HCPs) and candidate conservation agreements with assurances (CCAAs) (Langpap and Kerkvliet, 2012; Reeling, Palm-Forster, and Melstrom, 2017). Research in economics has also investigated the effectiveness of the ESA in

[1] There is also an earlier study by Freudenburg, Wilson, and O'Leary (1998) on the economic impacts of protecting public forest land for the northern spotted owl.

protecting imperiled species and preventing extinction (Metrick and Weitzman, 1996; Dawson and Shogren, 2001; Ferraro, McIntosh, and Ospina, 2007).

This paper estimates the employment impacts from listing an endangered species with habitat on private lands. We focus on the lesser prairie chicken, a grassland bird native to the southern Great Plains that was recently listed under the ESA. While the northern spotted owl threatened timber harvests on public land, the lesser prairie chicken threatens farming, ranching, and energy development. This case is not unique; the proposed listing of the greater sage-grouse is another recent example of ESA conflict on private lands (Smith et al., 2016). Our identification strategy takes advantage of the month listing occurred plus the spatial distribution of employment and the lesser prairie chicken's population. At the time of listing, individuals and firms had access to information on which privately owned lands were likely to be burdened by ESA regulations. Combined with panel data on county employment levels drawn from the BLS Quarterly Census of Employment and Wages, this information allows us to use quasi-experimental methods to test whether the number of jobs in counties with lesser prairie chickens declined because of the listing. We find evidence that employment changed by about 1.5% in counties with habitat, although the magnitude of the effect varies proportionally with the amount of habitat in a county. We also examine employment dynamics and the timing of conservation actions prior to the listing. We find evidence that conservation actions may have affected job growth before ESA regulations went into effect.

The Endangered Species Act

The ESA is Congress's attempt to prevent extinction events. The ESA, passed with bipartisan support in 1973, is the successor to several earlier laws, including the 1966 Endangered Species Protection Act and the 1969 Endangered Species Conservation Act. The 1966 Act was the first to authorize the Secretary of the Interior to develop a list of endangered species; those species received protection from habitat destruction on federal lands. The 1969 Act allowed the Secretary of the Interior to list foreign species and prohibited interstate commerce involving listed species or their products. However, a consensus soon emerged that these protections were insufficient. The Endangered Species Act of 1973 expanded listing categories to include endangered and threatened species and prohibited any act of harm to listed species, including those on private lands.[2] The U.S. Fish and Wildlife Service (FWS), the agency tasked with listing and protecting non-marine species, interprets the definition of "harm" broadly to include the destruction of species habitat. The Act further requires the FWS to designate critical habitat, which consists of areas essential to the conservation of a listed species. The ability to prohibit harm and designate critical habitat provides the FWS with powerful regulatory instruments for conservation.[3]

Today, the ESA is a controversial environmental law with support that tracks partisan lines (Brown and Shogren, 1998). This was not true at the time it was written—the law passed the Senate with a vote of 99 to 1—but several famous conflicts turned species listings into a contentious and high-stakes process. Just a few years after the ESA's passage, conflict erupted over the snail darter, a fish listed in 1975 because its range was restricted to a single section of one river. At the time, the Tennessee Valley Authority was completing a dam that would inundate and destroy the snail darter's

[2] The Act defines as "endangered" a species in danger of extinction throughout all or a significant portion of its range, while a species listed as "threatened" is likely to become endangered within the foreseeable future throughout all or a significant portion of its range.

[3] The law also uses Habitat Conservation Plans (HCPs), Safe Harbor Agreements (SHAs), and Candidate Conservation Agreements with Assurances (CCAAs) to incentivize private conservation. These instruments provide regulatory assurances to landowners vis-à-vis land use restrictions in return for practices that provide a conservation benefit for a listed or candidate endangered species.

habitat. The conflict culminated in a lawsuit widely covered in the media as a "classic struggle between ecology and economics" that eventually reached the U.S. Supreme Court (Plater, 2013).[4]

A similar controversy exploded in 1990 over the listing of the northern spotted owl, which resides in old-growth forests in the U.S. Northwest that serve as an important stock for the timber industry. Studies, some funded by trade groups, predicted that protecting the owl would cost tens of thousands of industry jobs (Goodstein, 1999). When factoring in the spillover to communities in the area dependent on logging and timber milling, the total effect was projected to reach hundreds of thousands of jobs (Meyer, 1997); subsequently, "jobs versus owls" became the slogan for anti-ESA politics. President George H. W. Bush famously commented, "We'll be up to our necks in owls, and every millworker will be out of a job." The President's remark was obvious hyperbole, but it testifies to public focus on job impacts as a critical measure of the costs of protecting endangered species.

The Lesser Prairie Chicken

The lesser prairie chicken is a species of long-standing concern. The grassland bird lives in parts of Colorado, Kansas, New Mexico, Oklahoma, and Texas, much of which are dominated by agriculture. By the 1990s, conversion to cropland and intensive grazing practices had reduced and fragmented the prairie chicken's habitat to 17% of its historical range, with population declines of up to 90% (Van Pelt et al., 2013). In 1995, the FWS received a petition to list the species. The agency determined a listing was warranted but delayed acting on it because resources were focused on higher priority species. However, emerging energy development in the habitat region prompted the agency to issue a proposal to list the species as threatened in December 2012. The lesser prairie chicken has a strong aversion to vertical structures, probably as an instinctual defense against birds of prey, so wind towers and oil and gas wells can be extremely disruptive (Haufler, Davis, and Caufield, 2012).

In response to increasing habitat threats and the proposed ESA listing, the Western Association of Fish and Wildlife Agencies (WAFWA) developed the Range-Wide Plan (RWP) (Van Pelt et al., 2013). The cornerstone of the plan is a conservation program that offsets habitat losses with new habitat brokered through voluntary land use agreements. Funding for these agreements comes from mitigation fees that developers pay to participate in the RWP, so that their projects qualify for the ESA's 4(d) rule, which exempts participants from the legal ramifications of "taking" a listed animal as long as the take is part of a plan that has a net benefit for conservation of the species. By "developers," we refer to individuals and companies that build on land. Many energy companies in the region lease rather than own land, but their activities are still subject to ESA regulations in the event of a listing. By participating in the RWP, developers can significantly reduce the risk of litigation from a take. The RWP was implemented soon after the FWS announced in May 2013 that exceptions would be allowed under the 4(d) rule if the lesser prairie chicken were listed (which at the time was still uncertain).

Many stakeholders expected that WAFWA's RWP would convince the FWS that a listing was unnecessary to protect the prairie chicken from further habitat losses. The FWS officially endorsed the plan and in December 2013 published a revised listing rule to clarify how the RWP would fit in the exceptions permitted by the ESA. However, in March 2014, the agency determined that the lesser prairie chicken should receive threatened species status, which was officially conferred in May 2014. This decision came as a surprise to many who had believed that the RWP would convince the FWS that listing was unnecessary.

The decision to list was widely criticized by industry (Steinhauer, 2014). Within a month, there were reports that the decision was having an effect on drilling decisions and energy jobs (Perry, 2014). Developers and politicians argued that the threatened species status would hinder

[4] The Supreme Court ruled in favor of the ESA, but its decision prompted Congress to exempt the dam from ESA restrictions. Biologists relocated a part of the snail darter population, which likely saved the species from extinction as the original population was extirpated when the dam was completed.

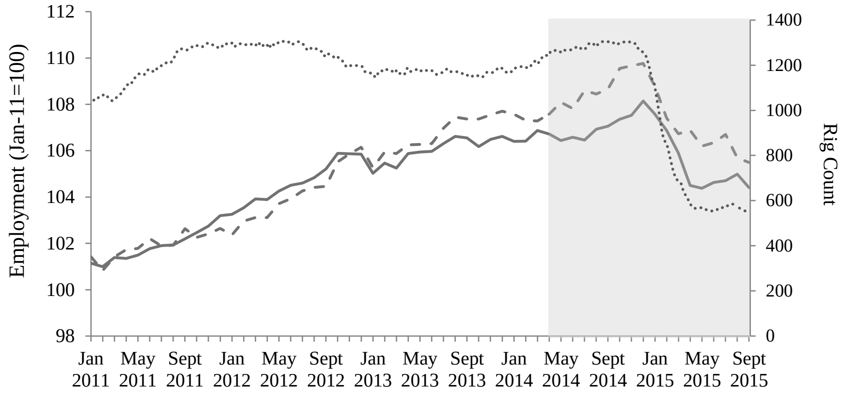

Figure 1. Employment Growth in Habitat Counties (solid line) and Comparison Counties (dashed line)

Notes: Seasonally adjusted and indexed to January 2011. The dotted line indicates rig count in the five-state area as a measure of regional drilling activity.
Source: Baker Hughes Rig Counts.

economic development in rural areas with habitat. One petroleum industry group publicly stated that ESA "regulations would impede operations and cost hundreds of millions of dollars in oil and gas development in one of the country's most prolific basins," while a U.S. Representative argued, "As the American economy continues to struggle, our actions should encourage growth not hinder economic efforts" (Associated Press, 2015). Several lawsuits challenged the listing decision, including one that resulted in the listing being overturned by a Texas federal judge in September 2015 (Wertz, 2015).

ESA regulations, or even the threat of regulations, can affect employment by reducing the expected net benefits of economic activity. The fact that some developers voluntarily participate in costly conservation programs is evidence that there are economic benefits to avoiding ESA regulations. In 2014, WAFWA received about $40 million in enrollment fees from the RWP. Lawsuits against individuals or companies for ESA violations are rare (so informed employers probably recognize that the probability of litigation is small), but the enrollment fees collected from the RWP suggest that the economic damages from ESA regulations are substantial. Rational employers will respond to regulatory costs by adjusting their investment and hiring decisions.

Graphical analysis suggests a shift in employment growth did occur following the listing of the lesser prairie chicken. Figure 1 presents the employment time series between 2011 and 2014 in counties with lesser prairie chicken habitat (solid line). In the period covered by our analysis, employment in counties with habitat was increasing by about 2% annually prior to May 2014. However, employment growth slowed in the second half of 2014. The figure shows a modest downward shift in the trend after the species was listed. Of course, the figure does not prove causality, but it provides contextual information that is consistent with the idea that ESA regulations can influence the labor market.

Empirical Strategy

We measure the local labor market effects of listing the lesser prairie chicken by comparing employment trends in counties with and without habitat. A decline in employment is expected in counties with lesser prairie chicken habitat following the listing. To test this empirically, we estimate a difference-in-differences model with the specification

$$\text{(1)} \qquad \ln(\mathrm{Y}_{it}) = \gamma_i + \tau_t + \delta\big(habitat_i \cdot listing_t\big) + \beta \mathrm{X}_{it} + \varepsilon_{it},$$

where Y_{it} is employment in county i in month t; γ_i are county fixed effects; τ_t are time-period (month) effects; $habitat_i$ is a measure of the habitat area; $listing_t$ is a dummy that takes the value of 1 if the month is in the interval May 2014 to September 2015, and 0 otherwise; and X_{it} are additional controls varying over location and time. The canonical difference-in-differences regression includes dummies for the treatment group and treatment time. Equation (1) is more general, with the treatment group and treatment time dummies absorbed by the county and period effects. We estimate equation (1) by OLS.

We also estimate an exponential specification,

$$\text{(2)} \qquad \mathrm{Y}_{it} = \exp\big(\gamma_i + \tau_t + \delta\big(habitat_i \cdot listing_t\big) + \beta \mathrm{X}_{it}\big)\eta_{it},$$

using the Poisson pseudo-maximum-likelihood (PPML) estimator with two-way fixed effects. The exponential form is useful for two reasons: First, equation (1) provides an estimate of the treatment effect on log employment, but we are interested in the effect on the non-transformed outcome. Second, OLS is consistent on the log scale only under a specific heteroskedastic error distribution (Santos Silva and Tenreyro, 2006). While we expect heteroskedasticity in modeling rural employment, in which the errors attenuate with smaller employment levels, there are benefits to having an estimator that does not hinge on strict distributional assumptions. Unlike OLS, the PPML estimator does not rely on heteroskedasticity assumptions for consistency as long as the conditional mean is correctly specified (Gourieroux, Monfort, and Trognon, 1984). The dependent variable does not have to be Poisson distributed, nor does it need to be a count. For inference that does not rely on the Poisson variance assumption and is robust to arbitrary patterns of serial correlation, one can use a sandwich estimate of the standard errors (Wooldridge, 1999).[5]

Under a common trend assumption, $100 \times [\exp(\hat{\delta}) - 1]$ is the average proportional effect of the treatment on the treated.[6] This is an estimate of the percentage point change in employment due to the intervention. This average treatment effect on the treated is

$$\text{(3)} \qquad \widehat{ATT} = \frac{1}{N_T} \sum_{i \in S_T} \exp\big(\hat{\gamma}_i + \hat{\tau}_t + \hat{\beta} \mathrm{X}_{it}\big)\big[\exp(\delta) - 1\big],$$

where S_T is the set of habitat counties and N_T is the number of such counties. For equation (3) to hold, the common trend assumption requires that the relative (not the absolute) change in employment is equal between groups prior to the intervention. This means employment must increase or decrease by the same percentage in the habitat and comparison county groups prior to the listing for $100 \times [\exp(\hat{\delta}) - 1]$ to be a credible estimate of the listing effect.

We consider two definitions of the treatment area $habitat_i$: First, $habitat_i$ is constructed as an indicator equal to 1 for counties with a positive amount of habitat, and 0 otherwise. The $100 \times [\exp(\hat{\delta}) - 1]$ thus becomes the difference-in-differences estimate of the proportional change in employment attributable to ESA regulations. Second, to exploit possible variation in the extent of ESA regulations across counties, $habitat_i$ is measured as the fraction of land designated as habitat in a county. In this case, $100 \times [\exp(\hat{\delta}) - 1]$ approximates the change in employment attributable to a pure (100%) habitat county. One would expect counties with more habitat to experience greater declines in employment if listing a species under the ESA causes a decline in local employment.

[5] See Santos Silva and Tenreyro (2010) and Limwattananon et al. (2015) for PPML applications with difference-in-differences.

[6] In the results, the estimated δ is small enough that $100 \times \delta$ is a good approximation (Wooldridge, 2012).

Data

We obtained information about the distribution of lesser prairie chickens from the Kansas Biological Survey, which has worked extensively with WAFWA to document areas of occupied and suitable habitat. These data are made available by the Southern Great Plains Crucial Habitat Assessment Tool (SGP CHAT), a spatial model that classifies habitat in the five-state region.[7] The SGP CHAT includes an online map showing priority habitat locations. The online interface was developed to provide public information and encourage development projects in sensitive areas to participate in the RWP, as the vast majority of habitat is on private land (Van Pelt et al., 2013). The SGP CHAT was published in 2013, so industry and the public had access to information about the distribution of the lesser prairie chicken at the time the species was listed.

Based on the SGP CHAT, 90 counties from a pool of 533 counties in the five-state region contain lesser prairie chicken habitat. Kansas contains the largest share, followed by Texas, Oklahoma, New Mexico, and Colorado. This allocation closely mimics the population distribution, with about half of all prairie chickens living in Kansas, followed by Oklahoma, Texas, New Mexico, and Colorado (Van Pelt et al., 2013). The amount of habitat ranges from 0.01% to 100% of county land, with an average of 62%.

For employment information, we use monthly county-level data from the Quarterly Census of Employment and Wages (QCEW). The U.S. Bureau of Labor Statistics conducts the QCEW using administrative data from employers who pay unemployment insurance premiums. The census includes monthly employment for every county in the United States. Employment is determined by place of work and measures total jobs—so a person holding multiple jobs is counted multiple times. The data do not include self-employed persons or farmers, ranchers, and military personnel, although hired farm workers are included. We use employment in all industries as the dependent variable. The QCEW publishes industry-specific employment data but suppresses these data at aggregation levels that include only a few establishments in a county, which is a problem in our study area. We do not report industry-specific impacts due to data suppression.[8]

We use a caliper matching algorithm to construct a comparison group that is observationally similar to the treatment group. Difference-in-differences without matching has been shown to perform poorly in estimating treatment effects (Imbens and Wooldridge, 2009; Ferraro and Miranda, 2014). We use nearest-neighbor (1:1) Mahalanobois matching with a propensity score caliper to minimize pre-treatment differences between habitat and comparison counties. The caliper improves the matching between the groups by setting a tolerance for the quality of the matches (Ferraro and Miranda, 2017). Matching dramatically improves the balance across several observable county characteristics, including overall employment and the percentage of employment in the agricultural and mining sectors,[9] with standardized differences below tolerable limits (Table 1) (Rosenbaum and Rubin, 1985). The caliper trims off six counties with habitat that did not have a sufficiently close match. Several counties in the comparison group match with more than one habitat county because matching is performed with replacement. Comparison counties are therefore weighted in the regressions by their match frequency.

While comparison counties are slightly less populated than habitat counties, we find that the two groups have similar proportional growth rates. We empirically tested the common-trend assumption by measuring the differences in comparison and treatment groups pre-listing following Autor (2003). There were no statistically significant differences between the two groups for any month between January 2010 and January 2014, indicating that employment in habitat and comparison counties

[7] The SGP CHAT is available at http://kars.ku.edu/maps/sgpchat/.

[8] Goods-producing jobs, which include construction, agriculture, and natural resource extraction, are the most likely to suffer from regulations. Exploratory industry-level estimates from the unsuppressed data indicate that employment effects are concentrated in the goods-producing sector relative to services. These results are available upon request. We expect the effects are larger in specific industries such as agriculture and energy, but missing data are a serious problem at the industry level.

[9] These values set to 0 in cases of data suppression.

Table 1. Employment Means in Habitat Counties and Comparison Counties

Variable	Counties with No Habitat – All	Counties with No Habitat – Matched	Habitat Counties	Standardized Difference
Total employment	35,250	4,115	4,332	4.40
Percent employment in agricultural sector	1.8%	5.1%	5.3%	2.29
Percent employment in mining sector	3.0%	4.9%	4.6%	−4.23
Employment growth between Jan. 2011 and Jan. 2014	4.6%	5.5%	5.3%	−1.56
Counties	443	67	84	

Notes: The standardized difference is calculated as $100 \times (x_T - x_M)/((s_T^2 + s_R^2)/2)^{0.5}$. Rosenbaum and Rubin suggest that bias is problematic when a standardized difference is greater than 20 (Rosenbaum and Rubin, 1985; Ferraro and Miranda, 2017).

generally grew at the same rate prior to treatment. Differences become more pronounced before 2010. Because the more recent data provide a better match (and are of more interest), we test for a causal employment effect using the post-2009 QCEW employment data. The period of investigation runs from January 2010—several years before the species was listed—to September 2015—when the listing was vacated. Figure 1 provides graphical evidence of the parallel trend. The figure also clearly shows employment growth between the two groups diverges after the prairie chicken was listed in May 2014, with the comparison group continuing to grow at about 2% annually until January 2015. In 2015, economic activity and employment declined in both sets of counties due to the fall in energy prices (although the price fall began in 2014, it did not affect drilling activity until 2015, as shown by the rig count line in Figure 1).

Results

Figure 2 shows that employment level differences accelerated after ESA regulations were put in place. Average county employment in the habitat group is always greater than the average in the set of comparison counties and was increasing comparatively faster prior to the listing date (the increase in *level* differences is naturally required to satisfy the common trend assumption in *log* differences for the regression analysis). In the months following the listing, though, this pattern reverses. The figure shows about 50 jobs lost per county in the habitat region a year after the listing, despite a pre-treatment trend of about 10 jobs gained relative to comparison counties. Together, these amounts provide a rudimentary difference-in-differences estimate of about 60 jobs lost due to ESA regulations. This estimate may be biased upward or downward to the extent that it ignores employment factors correlated with the treatment designation. Yet, as will be seen, after controlling for potential confounders through regression, the employment loss we estimate is not far off from the amount suggested in the figure.

Primary Results

Our regression estimates indicate that ESA regulations negatively affect employment. Estimates of equations (1) and (2) without any covariates (X) are shown in Table 2. Each cell presents an estimate of δ, depending on the habitat definition and estimator. The first row contains the OLS estimates and the second row contains the PPML estimates. Proceeding across the first row, the coefficient of -0.007 in column 1 indicates that employment changed by a relative -0.7% according to the log-linear model. Average employment in a habitat county was 4,332 in the pre-treatment period, so this estimate implies an average loss of 30 jobs per county. Column 2 presents the same result except the regression includes a habitat-specific linear trend (i.e., $habitat \times t$), which functions in the same manner as state-specific trends in difference-in-differences models that measure the effect

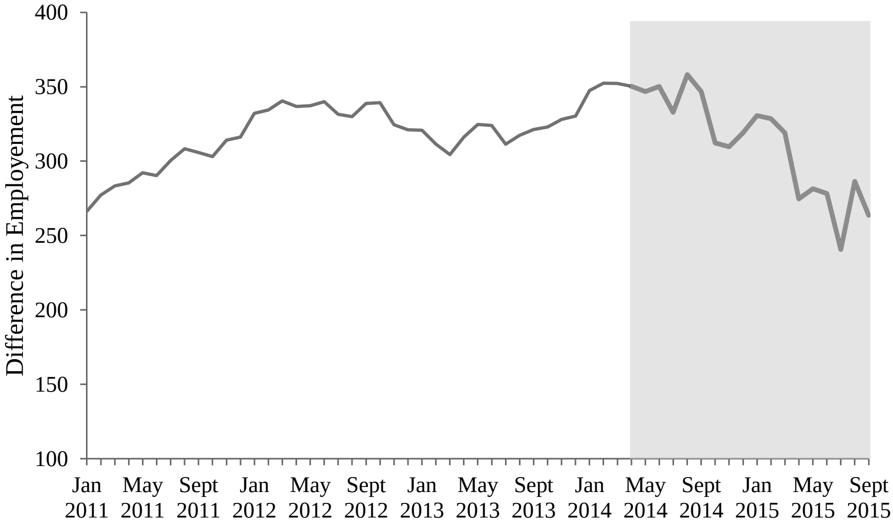

Figure 2. Difference in Average, Seasonally-Adjusted County Employment between Habitat Counties and Comparison Counties

Table 2. The Effect of ESA Regulations on Employment in Counties with Lesser Prairie Chicken Habitat. Each Cell Presents an Estimate of the Treatment Effect

Habitat Variable	**Post-Listing County Habitat Indicator**		**Post-Listing County Habitat Fraction**	
Estimator	**1**	**2**	**3**	**4**
OLS	−0.007	−0.027*	−0.015	−0.040***
	(0.014)	(0.016)	(0.011)	(0.014)
PPML	0.000	−0.022	−0.018	−0.046***
	(0.014)	(0.017)	(0.013)	(0.012)
Habitat-specific trend		X	X	

Notes: Single, double, and triple asterisks (*, **, ***) denote significance at the 10%, 5% and 1% level, respectively. Standard errors adjusted for clustering by commuting zone are reported in parentheses. The unit of observation is a county in a month. All models include county and period effects. The number of observations is 11,592.

of state policies. Including this trend increases the coefficient of interest to −0.027 in the log-linear model.[10]

By taking advantage of variation in the percentage of county land in habitat, we pick up stronger evidence of an employment decline. Estimates from comparable models using the fraction of land in habitat as the treatment measure are reported in columns 3 and 4 of Table 2. Based on the model estimated by OLS, we can say that for counties completely covered in habitat, employment changes by approximately −1.5% due to regulations. The effect is −4.0% when the trend is included.

The second row reports the results from the PPML estimator. For the model using the habitat indicator and without the habitat-specific trend, the PPML coefficient is a precise 0. The coefficient is −0.022 when the trend is added. When the habitat definition is changed to the fraction of land in

[10] We report standard errors clustered by commuting zone to allow for arbitrary cross-county and year correlations within commuting zones. Commuting zones were developed by the USDA Economic Research Service as a way to delineate local economies, typically containing ten or fewer contiguous counties.

habitat, the estimates are −0.018 without the trend and −0.046 with the trend, which are very close to their OLS-estimated counterparts.

Our preferred model includes the habitat-specific trend estimated by PPML. While the differences between the OLS and PPML-estimated coefficients are not large, we prefer the latter estimator because regression diagnostics suggest that the exponential specification of the PPML estimator is more appropriate. A Park-type regression test for heteroskedasticity (Park, 1966; Manning and Mullahy, 2001) does not support the conditional variance assumption of the log-linear model. The regression test estimated $\ln(\mathrm{Y}_{it} - \hat{\mathrm{Y}}_{it})^2 = \alpha + \beta \ln \hat{\mathrm{Y}}_{it} + v_{it}$, where $\hat{\mathrm{Y}}_{it}$ are the fitted values of Y_{it}. The OLS estimator of the log-linear model provides valid information about Y_{it} under the condition $\beta = 2$. We found $\beta = 1.6$ (p-value < 0.001 for a test of $\beta = 2$), so the OLS estimates are modestly biased. We therefore focus on the PPML results but note that, in general, both estimators provide evidence that ESA regulations affect employment.

Additional Controls

We next test the robustness of the results by adding variables for drought, commodity prices, and location-specific unobservable transitory factors. The drought index is included because the lesser prairie chicken was listed at a time of extreme drought in the five-state region. Negative index values indicate that an area received less than average rainfall in a month. This variable allows us to test whether drought in habitat counties drove the decline in employment observed after the ESA listing. Oil and gas prices interacted with the fraction of employment in the mining sector are included to control for their influence in counties that disproportionately rely on these commodities for economic development.[11] We also include the effect of wheat prices by interacting wheat prices with the fraction of employment in the agricultural sector in a county. It is possible that the effect of ESA regulations is confounded by declines in important commodity prices in states with relatively more habitat (i.e., Kansas, Oklahoma, and Texas). We do not include controls such as income, education, age distribution, etc., which are estimated for counties from annual surveys.

Overall, we find that the effects reported in Table 2 are largely insensitive to additional controls. The revised estimates are presented in Table 3, which shows that controlling for drought and commodity prices reduces the coefficient of interest. The estimate drops from −0.022 to −0.011 when habitat is measured as an indicator and from −0.046 to −0.042 when it is measured as a fraction. While the former coefficient is statistically insignificant, the latter is significant at the 5% level, which we interpret to mean that changes in drought severity, wheat prices, oil prices, and natural gas prices are not driving the results.

Columns 2 and 5 in Table 3 report the coefficients when state–period effects are added to control for arbitrary time-dependent factors (such as a common trend) affecting counties within each state. Adding this richer set of controls results in a slightly larger and more precisely estimated treatment effect, which is statistically significant at the 10% level in the specification that uses the county habitat indicator.

Columns 3 and 6 in Table 3 report the coefficients when county-specific trends are added to the model. These additional variables relax the assumption that employment trends are the same across counties within the habitat and comparison groups. Now, the effect of ESA regulations is identified by comparing employment changes before and after the lesser prairie chicken was listed after removing employment trends in each county, employment shocks for each state–year combination, and employment effects due to weather and prices. For these reasons, our preferred specification comes from column 3. In this model, the coefficient is −0.015, which is statistically significant at the 10% level. In column 6, the coefficient of interest is −0.028, which is statistically

[11] The value of this interaction equals 0 in cases where the BLS suppressed industry employment data. We would like to note that the effects of interest are not significantly affected by specifications that control for oil prices. Oil prices did not drop substantially until well after listing, in October 2014. The results are robust to narrowing the sample to the period prior to October 2014.

Table 3. Effect of ESA Regulations, Controlling for Drought, Commodity Prices, and Unobservable Transitory Factors Specific to States and Counties

	Post-Listing County Habitat Indicator			Post-Listing County Habitat Fraction		
Habitat Variable	**1**	**2**	**3**	**4**	**5**	**6**
ESA regulations	−0.011	−0.014*	−0.015*	−0.042***	−0.032*	−0.028***
	(0.011)	(0.009)	(0.009)	(0.013)	(0.019)	(0.009)
Palmer drought index	0.002	−0.002	−0.001	0.002	−0.002	−0.001
	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)	(0.001)
Wheat price	−0.019	−0.018	0.001	−0.019	−0.018	0.001
	(0.012)	(0.012)	(0.010)	(0.010)	(0.009)	(0.004)
Oil price	0.005***	0.005***	0.005***	0.005***	0.005***	0.005***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Natural gas price	0.035	0.025	−0.035**	0.034	0.024	−0.035**
	(0.025)	(0.027)	(0.015)	(0.024)	(0.027)	(0.014)
Habitat-specific trend	X	X		X	X	
State–period effects		X	X		X	X
County trends			X			X
County–month-of-year effects			X			X

Notes: Single, double, and triple asterisks (*, **, ***) denote significance at the 10%, 5% and 1% level, respectively. Standard errors adjusted for clustering by commuting zone are listed below the coefficients in parentheses. Oil price is denominated in $100s per barrel. All models include county and period effects.

significant at the 5% level.[12] This sizable difference indicates that counties entirely covered in habitat experienced greater than average employment losses.

Timing of Employment Changes

We now estimate the treatment effect with several monthly leads and lags to investigate the timing of employment changes with respect to ESA regulations. This specification interacts the treatment with dummy variables for each month running 6 months before the listing up to the month the listing was vacated (November 2013 to September 2015).[13] This allows us to examine how the employment trend in habitat counties differed from comparison counties a half-year prior to the listing. Employers may have anticipated a listing because the FWS made several pre-listing announcements about the status of the lesser prairie chicken. Most notably, the FWS revised its proposal to list the prairie chicken with the 4(d) rule in December 2013 to encourage participation in the RWP.

The estimated leads and lags from the model are plotted in Figure 3, which provides some evidence that employers anticipated a listing, responded to pre-listing conservation actions, or both. The first leading estimates suggest no difference between the habitat and comparison counties in terms of employment growth. Including additional leads does not change the interpretation of the figure, as they are close to 0. Although not shown here, we find no evidence of employment changes following any major FWS announcements prior to November 2013. In contrast, a small decline occurs after the December 2013 announcement and subsequent enrollment period in the RWP, with no appearance of a recovery in the following few months. Finally, a substantial and persistent decline occurs after ESA regulations went into effect.

[12] Adding county-by-month effects to control for county-specific seasonal employment patterns yields −0.027 (standard error = 0.010) for the coefficient of interest, which is statistically significant at the 5% level.

[13] The regression equation is $Y_{it} = \exp\left[\gamma_i + \tau_t + \sum_{\tau=-6}^{19} \delta_\tau \left(habitat_i \cdot \Phi_\tau\right) + \beta X_{it}\right] + \eta_{it}$, where Φ_τ are indicator variables for period τ, with $\tau = 0$ in the month of listing.

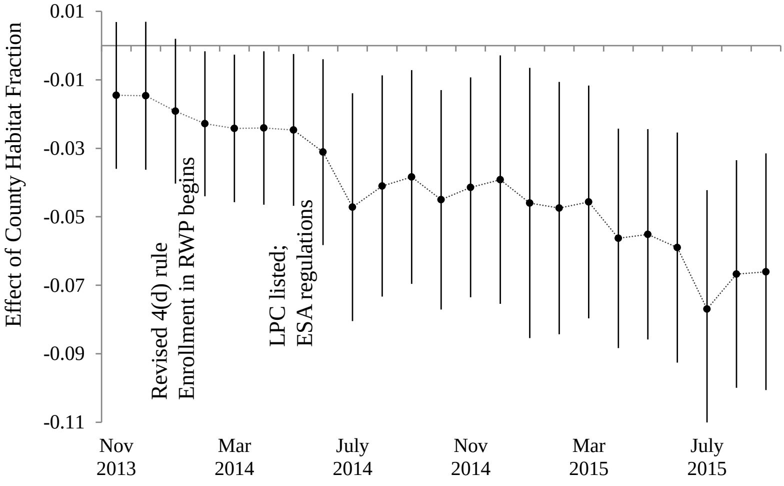

Figure 3. Estimated Employment Effect of the Fraction of Land in Habitat in the Months Before and After the Lesser Prairie Chicken Was Listed as Threatened

Notes: Vertical bars show 95% confidence intervals.

Table 4. Effect of ESA Regulations and Pre-Listing Announcements

Habitat Variable	Post-Listing County Habitat Indicator		Post-Listing County Habitat Fraction	
Variable	**1**	**2**	**3**	**4**
Revised 4(d) rule and beginning of RWP enrollment	−0.001	−0.013	−0.021*	−0.021
	(0.008)	(0.008)	(0.012)	(0.009)
ESA regulations	−0.018*	−0.008	−0.030***	−0.016***
	(0.010)	(0.007)	(0.008)	(0.009)
Habitat-specific trend	X		X	
Additional controls		X		X
State–period effects		X		X
County-specific trends		X		X

Notes: Single, double, and triple asterisks (*, **, ***) denote significance at the 10%, 5% and 1% level, respectively. Standard errors adjusted for clustering by commuting zone are listed below the coefficients in parentheses. All models include county and period effects. Additional controls include the drought index and commodity prices.

To statistically measure the effect of the RWP announcements, we estimated the model with habitat interacted with an indicator for the period following the FWS's announcement about the revised 4(d) rule and the start of the RWP. The results are shown in Table 4. Column 1 provides no evidence that the RWP announcements had an employment effect, while the effect of ESA regulations is significant at the 10% level. However, when state–period effects and county trends are added, the effect of the RWP and the effect of ESA regulations are negative and jointly significant at the 5% level. Columns 3 and 4 repeat these regressions except with habitat measured as a fraction. Significant at the 10% level, the coefficient on the December announcement indicates that a decline occurred prior to ESA regulations.

These results provide some limited evidence that employment in habitat counties declined prior to the ESA listing and that this decline occurred when the RWP was endorsed by the FWS and began enrolling developers in habitat impact mitigation agreements. This may indicate that developer agreements made through the RWP encouraged private conservation actions at the cost of established land uses, which subsequently affected employment levels. Unfortunately, we cannot determine with these data whether employment was influenced by habitat conservation actions. It is also possible that general awareness of the lesser prairie chicken's status fueled speculation among some employers that a listing would eventually occur.

Alternative Comparison Groups

Spillovers across counties could mean that our basic results misstate the impact of ESA regulations because the set of comparison counties includes some that border habitat counties. Employment effects will be overestimated if economic activities that would have occurred in habitat without the ESA listing move outside the habitat area, while effects will be underestimated if regulations also discourage economic activity adjacent to the habitat area. We may estimate a more accurate treatment effect by eliminating counties in the comparison group that buffer the habitat area.

Another potential concern with the benchmark comparison group is that many of the comparison counties at one time supported lesser prairie chicken habitat. The fact that these counties no longer provide suitable habitat suggests they may not be appropriate controls for the counties that do. Put differently, latent factors may be driving habitat loss and employment growth, and counties that no longer support habitat may be experiencing different employment growth patterns. Naturally, counties that buffer the habitat area also tend to have historically contained some habitat.

To address these concerns, we re-estimated the benchmark regressions after excluding counties that historically harbored lesser prairie chicken habitat from the set of comparison counties. Table 5 reports estimated treatment effects. For robustness, we also report the estimates from specifications that include state–period effects and county-specific trends. To make comparing the modified estimates easier, the first row reproduces estimates identical to those in Table 2 (columns 2 and 4) and Table 3 (columns 3 and 6). Omitting ex-habitat counties results in the loss of several thousand observations, but the estimates hardly change and thus provide no support for the hypothesis that ignoring development spillovers would lead us to overestimate the effect on employment. The effect changes from -0.015 to -0.016 using the modified set of comparison counties, which is only significant at the 15% level in the specification with the state–period effects and county-specific trends. When habitat is measured as a fraction of county land, the ESA effect remains negative and statistically significant at the 5% level (columns 3 and 4). Overall, these estimates suggest that the basic results are not biased due to spatial spillovers.

Conclusion

This paper presents evidence that ESA regulations negatively affect employment in areas with listed species. Applying a quasi-experimental method with panel data on employment, we found that counties with more habitat tend to suffer larger employment effects than those with less habitat. The precise estimate of this effect was somewhat sensitive to the type of estimator we used—in this case, OLS and PPML—but it was always negative.

There is some evidence that pre-listing conservation actions affected employment. Conservation agreements between private developers and wildlife agencies may be designed to reduce the regulatory implications of working on land with an endangered species, but these programs still have an economic cost. However, if declines in employment are attributable to participation in conservation agreements, the private sector is responding to conservation incentives. It is also possible that announcements about conservation actions led employers to anticipate a listing. In that case, changes in employment may have been temporary and would have returned to normal if the species had not been listed. This question deserves further study, as many ESA candidate species never receive threatened or endangered status.

Table 5. Effect of ESA Regulations

Habitat Variable	Post-Listing County Habitat Indicator		Post-Listing County Habitat Fraction	
Comparison Group	**1**	**2**	**3**	**4**
Benchmark	−0.022	−0.015*	−0.046***	−0.028***
	(0.017)	(0.009)	(0.012)	(0.009)
Counties with no historic habitat, away from habitat region	−0.021	−0.016	−0.049***	−0.030***
	(0.015)	(0.010)	(0.013)	(0.010)
Habitat-specific trend	X		X	
Additional controls		X		X
State–period effects		X		X
County-specific trends		X		X

Notes: The comparison group does not include counties that lost habitat due to development; most counties bordering the treated area are excluded in this robustness check. Each cell presents an estimate of the treatment effect. Single, double, and triple asterisks (*, **, ***) denote significance at the 10%, 5% and 1% level, respectively. Standard errors adjusted for clustering by commuting zone are listed below the coefficients in parentheses. All models include county and period effects. The additional controls includes the drought index and commodity prices. The number of observations is 11,592 for the benchmark sample and 9,177 with the restricted comparison group.

In our application to the lesser prairie chicken, we found that counties with habitat experienced a relative employment loss of about 1.5% when ESA regulations took effect. Average employment in affected counties was 4,332 before the listing, so this percentage decline implies a loss of 65 jobs per affected county. Overall, between 5 and 6 thousand jobs were lost due to the listing. Prior research estimates job losses due to protections for other species in the tens of thousands, so the effect we measure is comparatively modest. Furthermore, our estimate is a relative measure, and it appears that employment generally increased in the habitat area for some time after regulations were in place, albeit at a slower pace. At the same time, it is a real economic cost to lose thousands of jobs, especially when those jobs are located in areas with a dearth of local alternatives. We see the evidence in this paper as contributing important empirical data points to the debate on the economic costs of endangered species protection, although both sides of the conservation-versus-jobs debate may argue that the results here support their side.

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