Protected Area Effectiveness in European Russia: A Postmatching Panel Data Analysis

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ABSTRACT. *We estimate the impact of strict and multiple-use protected areas on forest disturbance in European Russia between 1985 and 2010. We construct a spatial panel dataset that includes five periods of change. We match protected areas to control observations and compare coefficients from fixed- versus random-effects models. We find that protected areas have few statistically significant impacts on disturbance, with little difference across parks closer to or farther from major cities or roads. Random-effects estimates differ qualitatively and quantitatively from those of fixed effects in our study, serving as a cautionary note for evaluations where time-invariant unobservables are important.* (JEL C14, Q23)

I. INTRODUCTION

Protected areas are a cornerstone for biodiversity conservation and the provision of ecosystem services such as carbon sequestration (Rodrigues et al. 2004; Scharlemann et al. 2010). They cover about 13% of terrestrial land, with continuing efforts to increase this area (Brooks et al. 2004; Jenkins and Joppa 2009). However, protected areas face many threats in conserving biodiversity and provisioning ecosystem services: they are often inadequately funded and staffed (Bruner et al. 2001) and are increasingly called on to meet multiple social objectives (Dudley et al. 1999; Naughton-Treves, Holland, and Brandon 2005; Sims 2010; Ferraro and Hanauer 2011). Furthermore, it is not always clear how much of an effect protected areas have, even in limiting forest loss, because where protected areas are placed strongly affects the additional benefits they bring to protecting biodiversity and ecosystem services (Joppa and Pfaff 2010). The majority of studies that examine protected area effectiveness have focused on the tropics, in particular Costa Rica and Brazil (e.g., Andam et al. 2008; Pfaff et al. 2009, 2013; Ferraro and Hanauer 2011; Nelson and Chomitz 2011; Nolte et al. 2013).

This paper's first objective is to estimate how effective different types of protected areas were at limiting forest disturbance in post-Soviet European Russia. The collapse of the Soviet Union in 1991 was one of the most dramatic political and socioeconomic changes in recent history, leading to rapid and unprecedented land use changes, including agricultural abandonment, decreased commercial logging, and increased illegal logging (Eikeland, Eythorsson and Ivanova 2004; Torniainen, Saastamoinen and Petrov 2006; Prishchepov et al. 2012). Little is known about how effective protected areas were during this time of rapid change. During this period, forest management in Russia changed several times, leading to confusion over management responsibilities (Sobolev et al. 1995; Colwell et al. 1997; Pryde 1997; Ostergren and Jacques 2002). There were also rapid decreases in budgets for biodiversity protection: one estimate puts posttransition budgets at about 10% of their 1989 levels (Wells and Williams 1998). During this same period, the number of protected areas expanded rapidly in Russia (Radeloff et al. 2013), and there are continued

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calls to increase the protected area network (Krever, Stishov, and Onufrenya 2009). Our analysis of the effects of protected areas on forest disturbance allows us to assess whether newly created parks have brought additional conservation benefit to European Russia, and to inform where new protected areas will contribute the most to protection of forests and, hence, biological diversity. Additionally, this information provides an important point for comparison to the recent studies of protected area effectiveness in tropical countries.

Establishment of protected forest areas is typically motivated by the desire to prevent some type of land use, development, or forest clearing activity that would be expected to occur in lieu of formal protection. However, there is a basic information asymmetry between those wanting to establish the protected area and the current user of the land. Only the current user/owner knows whether development or forest clearing would occur in the absence of protection. If minimal development or clearing would occur in the absence of protection, then the protected area generates only minor additional resource protection. However, if the land were to be cleared or developed in the absence of protection, then the protected area alters the land use outcome and generates additional benefits. Questions about such additional benefits are well-known problems with many environmental programs, such as carbon offset programs and the United Nation's REDD+ program to reduce deforestation (e.g., see Mason and Plantinga 2011).

To measure the additional impact of protected areas, or other environmental programs, requires careful attention to their nonrandom placement (Andam et al. 2008; Joppa and Pfaff 2009, 2010). Most protected areas are located in places unsuitable for other economic activities (Joppa and Pfaff 2009), and this remoteness reduces the impact that protected areas have on preventing logging or deforestation, because they have a lower probability of land cover change than areas outside of protection. For example, regressions that ignore the nonrandom placement of protected areas overestimate the impact of Costa Rican parks by as much as 65% (Andam et al. 2008). To control for the nonrandom placement of environmental programs, it is key to create a valid control group of observations that do not receive the program and use this group to estimate the impact of the treatment group (Ferraro and Pattanayak 2006; Ferraro 2009). Intuitively, if protected areas are located in remote areas unlikely to be developed, then the control parcels should be unprotected parcels that are also located in remote areas unlikely to be developed.

The second objective of our paper is to examine whether a combination of matching methods and panel data regression leads to different conclusions than matching with cross-sectional regression. Most recent evaluation studies of environmental programs use matching methods to construct a valid control group (e.g., Andam et al. 2008; Joppa and Pfaff 2011; Nelson and Chomitz 2011; Arriagada et al. 2012; Alix-Garcia, Kummerle, and Radeloff 2012; Alix-Garcia, Shapiro, and Sims 2012). The idea behind matching is to find the most similar observations to those that were protected, based on a selected set of covariates. Matching is typically combined with cross-sectional regression analysis to adjust for remaining differences in covariates. Since matching constructs a control group based on observables, omitted variables can still lead to bias in cross-sectional regression. To the extent that some omitted variables are time invariant (e.g., climate, soil quality), combining matching with fixed-effects panel regression methods provides an avenue to control for time-invariant plot-level omitted variables that can bias matching and or cross-sectional regression estimates (Cameron and Trivedi 2005). However, constructing panel data for impact estimates can be more costly and time-consuming than using cross-sectional data and may not even be feasible, depending on the date of program implementation. Thus, an important empirical question for the environmental program evaluation literature is whether investment in panel data is worth the effort, that is, whether fixed-effects estimates change conclusions relative to cross-sectional estimates.

A novel and distinguishing feature of this analysis is construction of a spatial panel dataset to analyze the impact of strict and multiple-use protected areas on forest disturbance in European Russia, a region that has received relatively little attention from land use researchers. We use satellite imagery to measure forest disturbance over five 5-year time periods from 1985 to 2010. We analyze the impact of six strict and six multiple-use protected areas created since the start of independence. Together, these 12 parks cover an area larger than the U.S. state of Rhode Island. By constructing a panel dataset we can combine matching—to control for bias arising from nonrandom placement of protected areas—with fixed-effects regression—to control for bias arising from time-invariant plot-level unobservables. This allows us to compare estimates from two methods that explicitly model a time-invariant unobserved plot effect (hereafter, plot effect) as either a fixed effect or a random effect. The plot effect includes any unobserved driver of forest disturbance that does not vary over the 1985–2010 period of our study (e.g., soil quality, climate, tree species). A random-effects approach to modeling the plot effect corrects estimated standard errors for serial correlation, but model identification rests on the assumption that the plot effect is uncorrelated with protected area status. This is unlikely if protected area locations are correlated with forest disturbance drivers (e.g., conservation of certain forest types). The random-effects identification assumption of no correlation between protected area status and the plot effect is implicit in most of the past literature measuring conservation effectiveness that uses cross-sectional data (e.g., Andam et al. 2008; Pfaff et al. 2009, 2013; Nelson and Chomitz 2011; Nolte et al. 2013). In contrast, fixed-effects modeling explicitly controls for the plot effect by de-meaning all model variables. Thus, identification with fixed effects no longer hinges on the assumption that the plot effect is uncorrelated with protected area status.

We find that strict protected areas reduce forest disturbance in European Russia by between 1 and 2 percentage points over some 5-year time periods but have no effect in others; indicating that overall, park status has little effect on disturbance rates. The impact of 1 to 2 percentage points may seem low, but is due to the fact that most strict protected areas are located far from threats and because the overall disturbance rates in European Russia during these five 5-year periods ranges from 2 to 6 percentage points. The magnitude of these effects is comparable to the global average impact of protected areas, which is about 2.5 percentage points (Joppa and Pfaff 2011). When we split the sample by distance to Moscow or major roads, we find little difference between parks located closer to or farther from threats. Multiple-use protected areas also have few statistically significant effects on reducing forest disturbance even though they are located in higher-threat areas. In stratifying multiple-use areas by distance to Moscow or major roads, we find limited evidence that parks closer to Moscow are more likely to experience higher rates of forest disturbance than comparable control observations.

Related to estimation strategy, we find evidence that fixed-effects estimates differ from random-effects estimates both qualitatively (i.e., statistical significance and sign) and quantitatively (i.e., magnitude of coefficient). These results are consistent with correlation between the unobserved plot effect and protected area status. The difference between fixed- and random-effects estimates varies by time period and threat level, and ranges from essentially zero to as much as a 60% difference in the coefficient on protected area effectiveness in our study. Since many conventional analyses of protected area effectiveness, as well as other environmental program evaluations, use matching with cross-sectional data, our results should serve as a cautionary note for analyses where the plot effect is important. To the extent that temporal variation in protected area status is available, our approach also highlights a potential solution to the identification problem arising from the plot effect being correlated with protected area status.

II. RUSSIAN FOREST MANAGEMENT AND PROTECTED AREAS SYSTEM

Forest Management

Following the collapse of the Soviet Union in 1991, the rates of forest disturbance decreased due to a mix of changing economic, social, and political conditions (Wendland et al. 2011; Baumann et al. 2012). In addition to the challenges of privatizing the timber industry, the forestry sector in Russia underwent multiple changes to management and governance that affected forest disturbance (Wendland, Lewis, and Alix-Garcia 2013). In the early 1990s, forest management and administration were decentralized to local and regional administrators, and the timber industry was privatized. The first official forestry legislation in post-Soviet Russia was the 1993 Principles of Forest Legislation. Under this legislation, the state maintained responsibility for forest management activities such as sanitary cuts, thinning, and reforestation, while former state logging enterprises and wood processing centers were privatized. Ownership of natural resources was excluded from privatization, but user rights, specifically the right to lease forests for industrial logging, were regulated in 1992 (Nysten-Haarala 2001). Leases for timber concessions could be short term (less than 5 years) or long term (up to 49 years).

In addition to changes to property rights, forest management and administration were initially decentralized to local forest administrators in 1993 (Krott et al. 2000; Eikeland, Eythorsson, and Ivanova 2004). Local forestry units operate on a scale roughly equivalent to administrative districts—equivalent to counties in the United States—in Russia. Poor forest management and inefficient utilization characterized these first few years of transition. These outcomes were largely due to the lack of technical skills and training provided to local state employees, and legislation that took away the primary source of funding for local forestry employees: timber harvesting. These changes in budgets created perverse incentives for local managers to charge high taxes and fees for timber contracts and to illegally cut timber to sell (Krott et al. 2000; Eikeland, Eythorsson, and Ivanova 2004; Torniainen, Saastamoinen, and Petrov 2006). These additional taxes and fees adversely affected the private timber industry. In addition, procuring markets for products and finding investment capital proved difficult for newly privatized firms (Pappila 1999; Kortelainen and Kotilainen 2003), leading to low rates of forest disturbance (Baumann et al. 2012).

In 1997, Russia issued its first Forest Code, which recentralized decision-making authority to the regional level—equivalent to states in the United States. This shift in authority away from local forest administrators helped reconcile the problem of high taxes and fees by making contracts between firms and the state more transparent. However, it failed to address the perverse incentives faced by local forestry units to cut timber illegally through the guise of sanitary logging in order to generate income (Torniainen, Saastamoinen, and Petrov 2006). During this period, economic conditions were also changing in Russia due to the end of a financial crisis and the devaluing of the ruble. There was a slight increase in forest disturbance over earlier post-Soviet periods during this time (Baumann et al. 2012).

In 2004, the central government recentralized forest authority, paralleling national shifts to regain control of regions. This is associated with a decrease in forest disturbance in Russia (Baumann et al. 2012). In 2007, Russia released its latest version of the Forest Code. This new Forest Code once again decentralized decision-making powers to the regional level and made the first substantive changes to forest property rights, designating several new responsibilities to firms and extending the duration of leases up to 99 years (Torniainen, Saastamoinen, and Petrov 2006).

Protected Areas

There are three types of federally protected areas in Russia: zapovedniks, national parks, and federal zakazniks. We group these into more generalizable categories: strict and multiple-use protected areas. Strict protected areas include Russia's zapovedniks. Zapovedniks are strict nature reserves, equivalent to an IUCN designation of Category I protected area, and logging and other extractive activities are prohibited (Wells and Williams 1998). The first zapovednik was established in the early 1900s, and at least a dozen new zapovedniks have been established in Russia since the collapse of the Soviet Union (Krever, Stishov, and Onufrenya 2009). Zapovedniks tend to be well funded and staffed compared to other types of protected areas; however, this financing is still inadequate to cover many of the costs of the parks (Wells and Williams 1998). Zapovedniks are managed by the Ministry of Environmental Protection and Natural Resources in Russia. Since there is no permitted logging within zapovedniks, evidence of logging within these protected areas is indicative of illegal activity.

We classify national parks and federal zakazniks as multiple-use protected areas. National parks are a fairly recent designation in Russia; the first national park was created in 1983, and more than a dozen have been created since the collapse of the Soviet Union (Krever, Stishov, and Onufrenya 2009). National parks were created to provide recreational and environmental education opportunities for people, and tend to be larger than other types of protected areas in Russia. They correspond to an IUCN Category II or IV protected area. There is designated federal funding for national parks; however, budgets vary considerably among parks. The Federal Forest Service managed national parks until 2000, which created several conflicts between intended and realized uses within the parks, since the primary mission of the Forest Service is industrial logging. Since 2000, national parks have been managed by the Ministry of Environmental Protection and Natural Resources (Ostergren and Jacques 2002). However, permits for logging within national parks are still granted on a case-by-case basis.

Federal zakazniks are one of the oldest forms of protection in Russia and correspond to an IUCN Category IV or V protected area. Several limited uses are allowed within federal zakazniks, such as grazing, hunting, and fishing. While there is no set management entity for federal zakazniks, the Ministry of Agriculture oversees many of them (Ostergren and Jacques 2002). Federal funding tends to be more limited for zakazniks compared to the other two types of federally protected areas, which impacts staffing and enforcement (Pryde 1997). It is difficult to determine whether logging is legal or illegal in a given zakaznik, since logging permits can be granted, but the lack of monitoring and enforcement also means that illegal logging is possible (Sobolev et al. 1995). Thus, for both types of multiple-use protected areas, evidence of forest disturbance could be indicative of logging permitted by the federal government, or illegal logging activity.

In sum, the widely varying changes to timber management and funding in both protected and unprotected areas from 1990 to 2010 require careful consideration in constructing an empirical analysis that allows for temporally heterogeneous treatment effects when comparing forest disturbance across protected and unprotected areas in this region.

III. STUDY AREA AND DATA

Study Area

Our study area includes 12 federally protected areas covering 4,045 km^2 in the temperate forest zone of European Russia: six strict and six multiple-use areas (Figure 1). This total area is about one-third the size of the protected areas system in Costa Rica (Pfaff et al. 2009) and slightly larger than the total land and water area of the U.S. state of Rhode Island. The average strict protected area in our sample is 191 km^2 in size; multiple-use areas tend to be larger, with an average size of 483 km^2. The date of establishment of these 12 protected areas varies between 1989 and 2006 (Table 1). While there were some protected areas in our study region established prior to the collapse of the Soviet Union, we analyze only post-Soviet protected areas, in order to provide a fair assessment of fixed- versus random-effects models.

Central European Russia is a mosaic of agriculture and forest. Agricultural crops include mostly grains, and the southern part of the study area includes the fertile "black soil" zone. The forests of Central Russia are made up of deciduous and coniferous tree species. Common deciduous species include lime, oak, birch, aspen, ash, maple, and elm. Scotch pine is the dominant coniferous species. While total forest cover is lower in this region than in parts of Northern European Russia, timber harvesting is still important due to low transportation costs. In particular, timber harvesting around the city of Moscow has in-

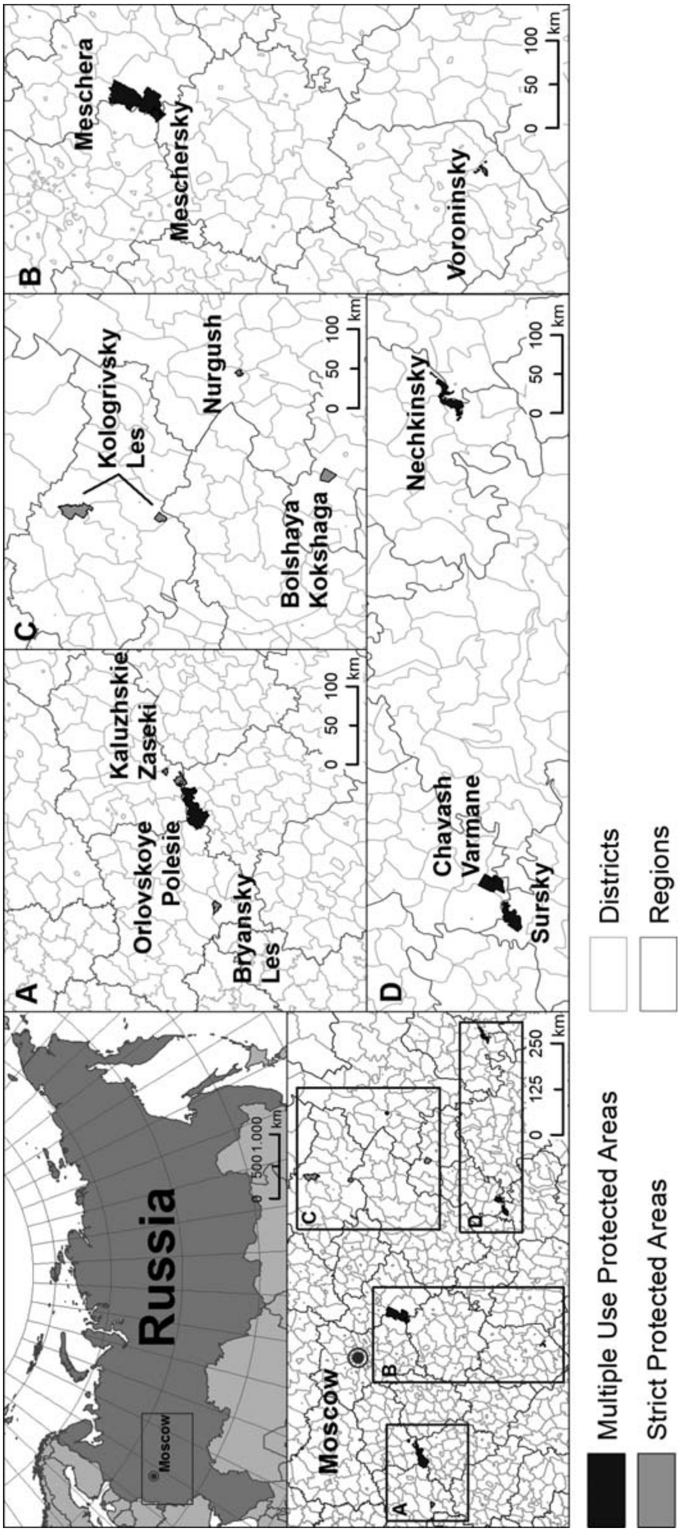

FIGURE 1
Locations and Names of Protected Areas

TABLE 1

Number of Observations and Forest Disturbance by Year and Protected Area Type in the Full Sample (Unmatched Sample)

	1985–1990	1990–1995	1995–2000	2000–2005	2005–2010
All Pixels Outside of Protected Areas					
Total pixels	215,477	203,307	192,334	186,889	179,212
Percent forest disturbance	5.65	5.40	2.83	4.11	2.18
Strict Protected Areas					
Total pixels		1,103	6,131	6,080	9,773
Percent forest disturbance		1.18	0.83	3.60	0.17
Parks		1	5	5	6
Park area (km^2)		97	557	557	1,147
Multiple-Use Protected Areas					
Total pixels		2,136	15,134	20,957	19,689
Percent forest disturbance		12.08	3.29	6.05	2.49
Parks		1	5	6	6
Park area (km^2)		238	2,674	2,897	2,897

creased considerably since 2000 (Wendland et al. 2011; Baumann et al. 2012). Population density in Central Russia is also higher than in other parts of Russia, potentially causing higher threats for protected areas.

Forest Disturbance Data

We define protected area effectiveness as the change in forest disturbance due to protected area designation. Forest disturbance is measured using remote sensing imagery. Forest cover is strongly correlated with biodiversity and is the outcome evaluated in most assessments of protected area effectiveness (e.g., Mas 2005; Andam et al. 2008; Pfaff et al. 2009, 2013; Joppa and Pfaff 2011). While we cannot distinguish forest disturbance events due to manmade versus natural causes—a limitation in most land use change analyses—remote sensing classifications of forest disturbance in European Russia have attributed the majority of disturbance events to manmade causes such as logging (Potapov, Turubanova, and Hansen 2011). Additionally, the remote sensing data used in this study were visually inspected, and any data for parks and years that were characteristic of a natural event such as flooding or fire were removed.

Our measure of forest disturbance comes from eight Landsat footprints classified for forest cover change in 5-year increments from 1985 to 2010 (see Baumann et al. 2012). This primary analysis provides 30 m resolution data on forest cover change, with average accuracies greater than 90%. We randomly sample 1% of all pixels within each of the 12 protected areas that were forested according to the 1985 land cover classification. Thus, we take an equal proportion of pixels from each park. This gives a sample size of about 36,000 protected area pixels. We then sample four times this amount of forested pixels from areas outside of protected areas. For both samples we specify a minimum distance criterion of 300 m between pixels to reduce spatial correlation.

For each pixel selected we record whether it stayed in forest in a given 5-year period (value of "0") or whether it was disturbed by forest clearing ("1"). A pixel is removed from the dataset once the forest is disturbed, because 20 years is not sufficient time for forest to regenerate to a harvestable size given an average rotation period of more than 50 years in Russia. Because pixels are removed once cut, and because new protected areas were created between 1990 and 2010, the total number of observations within and outside of protected areas varies over each time period (Table 1): in 1985–1990—before any parks—there are approximately 215,000 unprotected observations, whereas in 2005–2010 there are about 30,000 within protected areas and 180,000 outside of protected areas.

Covariates

We select covariates that we assumed to be correlated with both the treatment (protection) and outcome (forest disturbance), and that are available for our study region. In the tropics, protected area placement has been found to be highly correlated with remoteness and low economic productivity (Andam et al. 2008; Joppa and Pfaff 2010). Forest disturbance in European Russia has shown to increasingly be correlated with profit-maximization behaviors that factor in transportation costs and opportunity costs of the land (Wendland et al. 2011). Thus, for both of these decisions—protection and forest disturbance—we control for accessibility and biophysical characteristics of the pixel, as these characteristics strongly influence the net economic returns from disturbing a forest plot.

Forest disturbance is a capital-intensive activity whose net returns are greatly affected by accessibility. Since we lack monetized plot-level cost variables, we include multiple physical proxies of disturbance costs that are related to accessibility. Specific variables include the distances to forest edge, closest town, Moscow, and closest road; elevation; and slope. Distances are measured as the Euclidean distance from the pixel to the object and recorded in kilometers. Datasets on Russian cities with at least 50,000 persons, and major paved roads (circa 1990), are from the ScanEx Research and Development Center,[1] a Russian remote sensing company. Data on forest edge are derived from the remote sensing analysis described above, and calculated for each time period, whereas other distance measures did not vary over time. Elevation and slope data come from NOAA's Global Land 1 km Base Elevation Project;[2] elevation is measured in meters and slope as a percent.

There are additional biophysical variables that might be correlated with timber productivity, such as climate, soil quality, or rainfall. Since these physical accessibility indicators likely influence disturbance returns, the same indicators will also affect protected area status if regions with low returns to disturbance are systematically more (or less) likely to be protected than regions with high returns to disturbance. However, variables such as climate and soils are generally time invariant over the 25-year period of our study, and the fixed-effects estimation strategy (see Section IV) controls for these by placing all variables in difference-in-means form. As argued below, placing variables in difference-in-means form implicitly controls for all time-invariant forest disturbance drivers by eliminating them from the model unobservable in estimation. Since the net returns to forest disturbance can be strongly impacted by time-invariant physical plot characteristics, panel analysis with plot fixed effects provides a simple way to control for important drivers of forest disturbance without collecting additional data.

IV. EMPIRICAL STRATEGY

Our objective is to estimate the average treatment effect on the treated (i.e., protected areas), which is the difference between forest disturbance within protected areas and the expected effect if the protected area were not there. Mathematically, this is represented by

$$\tau = \frac{1}{N} \sum_{1, \mathbf{P}_i = 1}^{N} \mathbf{D}_i(1) - \mathbf{D}_i(0), \qquad [1]$$

where $\mathbf{P}_i = 1$ when a plot, i, is protected, and $\mathbf{D}_i(\cdot)$ is the observed outcome with "1" indicating forest disturbance and "0" otherwise. This gives the amount of forest disturbance prevented within the boundaries of the parks by protected area status. We estimate treatment effects separately for strict and multiple-use protected areas.

To construct a valid control group we use matching to select the best controls from observations outside of protected areas (Table 1). For each protected area type we partially control for administrative influences on forest clearing discussed in Section II by omitting any control observations that do not fall within the same administrative regions as the protected areas (see Figure 1). We then use

[1] See www.scanex.ru/en.

[2] See www.ngdc.noaa.gov/mgg/topo/globe.html.

logistic regression on the remaining observations to estimate the propensity score, that is, the conditional probability of a treatment (i.e., protected area observation) or control observation being designated as a park. Specifically, we estimate

$$\text{Prob}(\mathbf{P}_i = 1) = F(\boldsymbol{\alpha} + \varphi\mathbf{X}_i), \quad [2]$$

where $\mathbf{X}_i$ are the observable covariates described in Section III; and F is the logistic function.

The estimated propensity scores are then used to match treatment to control observations using nearest neighbor one-to-one matching without replacement, as suggested by Rubin (2006). To match the data, we estimate the propensity score (equation [2]) for each protected area type using the 1985 remote sensing data—before any of the protected areas were designated in our study. Thus, we assume that 1985 matched treatment-control observations remain good matches through all time periods. We implement matching using the PSMATCH2 algorithm in Stata11 (Leuven and Sianesi 2003).[3] We restrict the maximum distance between matches using a caliper size of a quarter of the standard deviation of the estimated propensity score, as recommended by Guo and Fraser (2010).

To ensure that matching improves similarity between treatment and control observations, we check covariate balance in our samples before and after matching by calculating the difference in means normalized by the square root of the sum of the variances, which is preferable over the t-statistic when there are large differences in sample size (Imbens and Wooldridge 2009). Specifically, we estimate

$$\bar{X}_1 - \bar{X}_2/\sqrt{\sigma_1^2 + \sigma_2^2}, \quad [3]$$

where $\bar{X}$ is the mean, σ^2 the variance, and "1" designates areas within protected areas and "2" areas outside of protected areas. The rule of thumb is that a normalized difference in means greater than 0.25 can bias regression estimation (Imbens and Wooldridge 2009).

We estimate postmatching linear regressions for each time period as

$$\mathbf{D}_{it} = \boldsymbol{\alpha} + \boldsymbol{\rho}\mathbf{Z}_i + \boldsymbol{\delta}\mathbf{P}_{it} + \boldsymbol{\gamma}\mathbf{YEAR}_t + \boldsymbol{\theta}\mathbf{P}_{it}\mathbf{YEAR}_t + \boldsymbol{\beta}\mathbf{Dist\_Edge}_{it} + \boldsymbol{\mu}_i + \boldsymbol{\varepsilon}_{it}, \quad [4]$$

where t indicates the time period (1985–1990, etc.); $\mathbf{D}_{it}$ is "1" if plot i is disturbed in time t and "0" otherwise; $\mathbf{Z}_i$ consists of the set of time-invariant independent variables (e.g., distance to Moscow) contained within the larger set of previously described covariates $\mathbf{X}_i$; time-varying independent variables are protected area status ($\mathbf{P}_{it}$) and distance to the forest edge ($\mathbf{Dist\_Edge}_{it}$); $\boldsymbol{\mu}_i$ are plot effects; and $\mathbf{YEAR}_t$ is a vector of year fixed effects used to control for variations over time that affect all observations (national timber prices, exchange rates, etc.). Estimated parameter vectors include the set $\{\boldsymbol{\alpha},\boldsymbol{\rho},\boldsymbol{\delta},\boldsymbol{\gamma},\boldsymbol{\theta},\boldsymbol{\beta}\}$. The parameter vectors $\boldsymbol{\delta}$ and $\boldsymbol{\theta}$ are used to form the average marginal effects of a protected area on the plot-level probability of forest disturbance accounting for the interactions between protection status and time dummies.

A key identification question is how to handle the plot effect $\boldsymbol{\mu}_i$, which includes all time-invariant plot-level omitted variables. These omitted variables can bias the parameter $\boldsymbol{\delta}$ if they are correlated with both the likelihood of being a protected area and forest disturbance. For example, we lack good data on a plot's soil quality and the microclimate in which the plot resides. Soils and microclimatic conditions can influence the type of tree species that can be grown and, hence, timber yields. Such unobservables are likely to be time invariant over the 25 years of our data and could influence protected area decisions if the government would like to conserve particular forest types. These unobservables can induce bias if the plot effect $\boldsymbol{\mu}_i$ is modeled as a random effect. However, modeling the plot effect as a fixed effect provides a way to control for any such time-invariant unobservable and observ-

[3] One limitation of propensity score matching is that standard errors are incorrectly estimated; this would lead to erroneous conclusions of the statistical significance of protected areas if the treatment effect was calculated directly from the matched data (i.e., through difference in means as implied by equation [1]). However, since we use matching to preprocess our data and restrict our sample before regression analysis, that is, we do not use the propensity score directly to estimate treatment effects, this does not affect our analysis.

able drivers of forest disturbance. By writing equation [4] in differences-in-means form (known as the within estimator), all time-invariant variables are eliminated (including $\boldsymbol{\mu}_i$), while all parameters on time-varying variables are preserved. Only parameters on time-varying independent variables can be identified when plot fixed effects are modeled; in our case these are only the distance to forest edge and the protected area dummy.

Our inclusion of an interaction between the park dummy variable and the vector of year fixed effects allows the marginal effect of protected area status to vary by time period, an important feature given the large temporal variation in policy factors affecting timber management and conservation budgets from 1990 to 2010 in European Russia. Additionally, we estimate equation [4] without the interaction term to generate an average park effect. We follow Pfaff et al. (2009, 2013) in testing for heterogeneity effects across parks by estimating equation [4] for parks above and below the median distances to Moscow and the nearest road, important physical indicators of forest disturbance threats.

We estimate equation [4] as a linear probability model in both random- and fixed-effects form. We choose a linear probability model over nonlinear probit or logit models for two reasons. First, plot fixed effects cannot be easily included in nonlinear discrete-choice models in a flexible manner. Plot fixed effects cannot be included in a probit model due to the incidental parameters problem, and fixed-effects logit models do not allow for calculation of marginal effects, since marginal effects are nonlinear functions of the unestimated fixed effects (see Wooldridge 2010, ch. 15). Further, while correlated random-effects estimation can be used in nonlinear models to introduce some correlation between a plot-level random effect and the park dummy variable (see Cameron and Trivedi 2005, ch. 23), one must assume a particular distribution for the plot effect. In contrast, the fixed-effects linear probability model is robust to any distributional assumption of the plot effect, as the plot effect is entirely differenced out of the model. Second, while linear probability models have the obvious weakness of not constraining probabilities between zero and one, many researchers have shown that they give almost identical marginal effects at the mean of the data, as do nonlinear probit or logit models with similar identifying assumptions (see Angrist and Pischke 2009, ch. 3; Wooldridge 2010, ch. 15). Since our primary interest is in estimating marginal effects, we choose the more flexible fixed-effects linear probability model for estimation. The functional form assumption in equation [4] is that treatment is linear and additive.

Finally, all estimations include standard errors clustered at the district level (see Figure 1). Cluster robust standard errors allow spatial correlation across units, in our case, correlation across pixels within the same district. The district level is important for forest management decisions (see Section II) and provides a reasonable spatial distance to allow correlation across units without imposing strong distributional assumptions on the data. Cluster robust standard errors also control for general forms of serial correlation (Cameron and Trivedi 2005).

V. RESULTS

A fundamental assumption in our fixed-effects strategy—commonly referred to as the parallel trends assumption—is that forest disturbance would be the same on control and protected plots in the absence of protection, and it is the protected area status that would induce deviation from a common underlying disturbance trend. We evaluate the parallel trends assumption with a temporal plot of the average forest disturbance probability in plots that are (ever) protected and on matched controls (Figure 2). Forest disturbance probabilities in our sample of pixels outside of protected areas range between 2 and 6 percentage points over a 5-year time period. Disturbance rates have generally fallen since the collapse of the Soviet Union, with an increase in disturbance in 2000–2005 that corresponds to the end of the Asian financial crisis (1998) and the beginning of Putin's presidency (2001). These temporal patterns in forest disturbance are consistent with reports on logging trends in post-Soviet Russia (Torniainen, Saastamoinen, and Petrov 2006) and remote sensing analyses of forest cover in European Russia

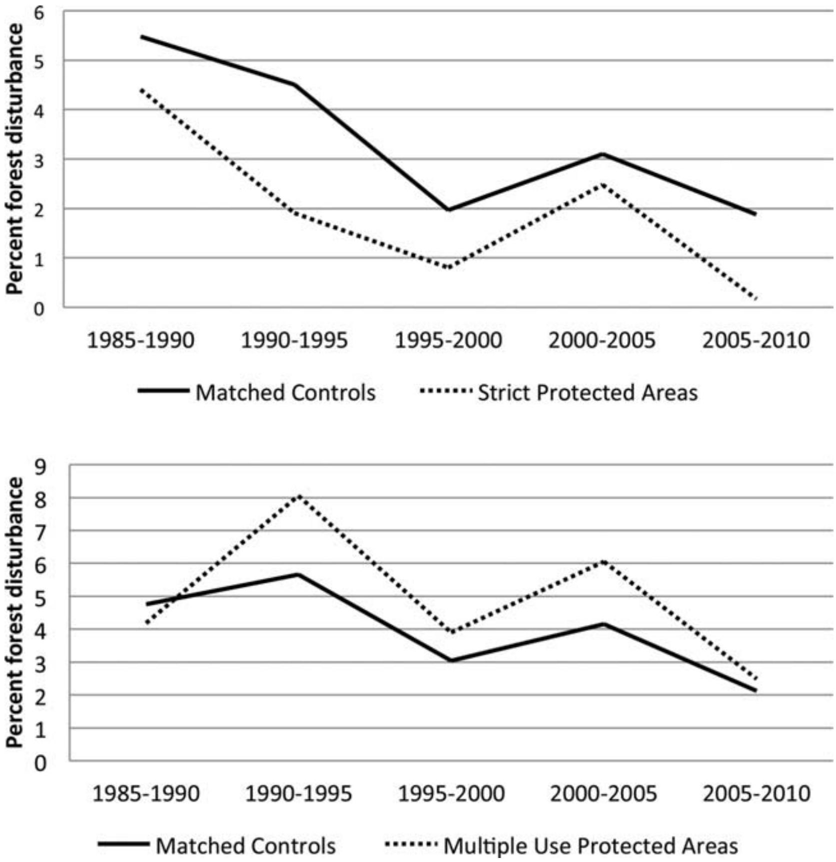

FIGURE 2
Forest Disturbance Trends in Protected and Control Plots: Strict Protected Areas and Matched Controls *(upper),* Multiple-Use Protected Areas and Matched Controls *(lower)*

(Potapov, Turubanova, and Hansen 2011; Baumann et al. 2012).

The similarity of forest disturbance trends across protected and control plots (Figure 2) is striking and strongly suggests that parallel trends is a reasonable assumption for our data. Rather than one single protected area event as in some difference-in-difference analyses, our study region is characterized by gradual increases in the protected area stock over the 1990s and 2000s (Table 1). Therefore, rather than one discrete drop in average disturbance probabilities for protected plots at the time of protection, we would expect to see a gradual divergence in forest disturbance on those plots that are (ever) protected compared to control plots if the presence of protection had a significant treatment effect. No divergence is visually apparent in Figure 2 between either strict protected areas and control plots, or between multiple-use areas and control plots. Therefore, there is no strong visual evidence of a significant negative treatment effect of protected area status on forest disturbance. A formal postmatching econometric analysis is needed to further evaluate this finding.

Considering where protected areas are located, descriptive statistics suggest differences across park types and controls (Table 2); each of these differences is statistically significant at the 1% level using conventional *t*-statistics. Strict protected areas are more likely to be farther from the forest edge and closest town and road than the average control

TABLE 2
Summary Statistics and Difference in Means for Protected Areas and Areas Outside of Protection

Variable	All Pixels Outside of Protected Areas	Strict Protected Areas	Multiple-Use Protected Areas
Distance to forest edge (km)	0.23 (0.28)	0.42*** (0.44)	0.19*** (0.21)
Distance to closest town (km)	74.45 (47.74)	113.65*** (60.76)	59.37*** (25.85)
Distance to Moscow (km)	443.62 (247,59)	484.49*** (125.69)	528.33*** (352.79)
Distance to closest road (km)	1.19 (1.06)	1.69*** (1.27)	1.50*** (1.28)
Elevation (m)	154.15 (40.77)	163.74*** (47.81)	136.46*** (47.13)
Slope (%)	1.27 (1.41)	1.17*** (1.26)	1.38*** (2.15)
Observations	215,477	10,775	24,752

Note: Standard deviations in parentheses. Summary statistics are based on total number of pixels sampled within protected areas and outside of protected areas (i.e., before matching) in 1985. If observation was ever a protected area (became a protected area in 1995, 2000, etc.) it was summarized in the protected area column. The distance to forest edge is reported for 1985 values. The difference in means between observations outside of protected areas and within protected areas is estimated using unequal t-tests. Statistical differences between that type of protected area and control observations are reported by asterisks in the protected area columns.
* $p<0.1$; ** $p<0.05$; *** $p<0.01$.

TABLE 3
Covariate Balance Using Normalized Difference in Means

	Strict Protected Areas vs. Controls		Multiple-Use Areas vs. Controls	
Variable	Unmatched	Matched	Unmatched	Matched
Distance to forest edge 1985	0.21	0.09	−0.11	−0.05
Distance to closest town	0.18	−0.04	−0.13	0.02
Distance to Moscow	−0.18	0.11	0.17	−0.10
Distance to closest road	0.33	0.04	0.25	−0.11
Elevation	−0.19	−0.05	−0.41	−0.11
Slope	−0.15	−0.03	−0.01	0.02

Note: Normalized differences in means are estimated as the difference in the mean values of the covariates across protected area types and their control groups, normalized by the square root of the sum of the two variances. A negative sign indicates a smaller value for the protected area and a positive sign indicates a larger value for the protected area. The rule of thumb is that linear regression methods tend to be sensitive to a normalized difference in mean greater than 0.25 (Imbens and Wooldridge 2009). Unmatched sample included all control observations (Table 2, column 2). Matched sample was based on one-to-one nearest-neighbor matching without replacement using a caliper.

observation, indicating remoteness. But, they are closer to Moscow and have lower elevations and less steep slopes than control observations—factors that could lead to more harvesting. Multiple-use areas tend to be farther from roads, farther from Moscow, and at lower elevations than observations outside of protected areas (Table 2). However, these parks are on average closer to the forest edge and nearest town than the average control observation, indicating that they are in closer proximity to logging threats.

A more formal test of differences among treatment groups and controls is the normalized difference in means (Table 3). For both strict and multiple-use protected areas, several covariates have differences in means exceeding the rule of thumb of 0.25, indicating that simple regression analysis without matching is likely biased (Imbens and Wooldridge

TABLE 4
Estimates of Protected Area Impact on Forest Disturbance Using Matched Sample

	Random Effects	Fixed Effects
Strict Protected Areas		
Overall park effect	0.15% (0.54)	− 0.82% (0.60)
1990–1995	− 1.24% (0.78)	− 1.09% (1.10)
1995–2000	0.53% (0.58)	− 0.72%* (0.42)
2000–2005	2.18% (1.72)	− 0.69% (1.86)
2005–2010	− 1.33%** (0.57)	− 1.89%*** (0.31)
Observations	100,668	100,668
Multiple-Use Protected Areas		
Overall park effect	0.47% (1.02)	1.48% (1.98)
1990–1995	3.32%*** (1.14)	4.05%*** (0.90)
1995–2000	− 1.42% (0.96)	− 0.69% (1.27)
2000–2005	2.15% (2.24)	3.13% (3.21)
2005–2010	− 0.08% (0.50)	1.53% (1.97)
Observations	184,185	184,185

Note: Standard errors in parentheses. Time period effects (1990–1995, etc.) are estimated from equation [4]. The overall park effect is estimated after dropping the interaction term from equation [4] and rerunning the regression. The estimated parameters are the average marginal impact of protected areas on forest disturbance during that time period. Cluster robust standard errors are used in all regressions to control for spatial and serial correlation at the plot level.
* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.

2009). After matching, the covariate balance is greatly improved for all covariates and both protected area types (Table 3). The slight remaining differences in means further motivate the use of postmatching multivariate regression analysis.

The impact estimates for strict protected areas suggest few statistically significant effects: there is a negative and statistically significant effect of about 2 percentage points in 2005–2010 and about 1 percentage point in 1995–2000, although the latter is only weakly statistically significant at the 10% level (Table 4). The negative sign indicates that strict protected areas experience less forest disturbance than comparable control observations in these time periods. The estimated effects are statistically significant in time periods when there is low forest disturbance in the study area (Table 1). Random- and fixed-effects estimators give reasonably similar results for most time periods, although there is a divergence in the qualitative and quantitative results for whether parks are effective or not in 1995–2000 and in the quantitative effect in 2005–2010. Estimated as an overall average treatment effect (i.e., dropping the time interaction in equation [4] and reestimating the model), we find no significant effect of strict protected areas on forest disturbance in post-Soviet European Russia.[4]

When we estimate the effect of strict protected areas based on their location, we find slight differences across parks located closer to or farther away from Moscow (Table 5). These results are reported only after 1995 since there is only one park in the 1990–1995 time period (see Table 1) in the sample; splitting the effects for one park across distance to Moscow or road did not seem relevant for policy implications. We find that parks closer to Moscow (higher pressure) reduce forest disturbance by about 2 percentage points in 1995–2000 and 2005–2010, but have as much as 4 percentage points more forest disturbance than similar plots outside of protection in 2000–2005. Parks located farther from Moscow (lower pressure) reduce forest distur-

[4] We also examined the heterogeneity of treatment effects by estimating a duration model where the effect of parks-since-treatment was modeled. We found statistically significant effects modeling the treatment effect this way but did not observe any generalizable conclusions about what length of time was needed after a park was established before it resulted in a treatment effect.

TABLE 5
Estimates of Protected Area Impact on Forest Disturbance Stratified by Distance to Moscow

	Random Effects		Fixed Effects	
	Lower Pressure (above median value)	Higher Pressure (below median value)	Lower Pressure (above median value)	Higher Pressure (below median value)
Strict Protected Areas				
Overall park effect	− 0.75% (0.78)	1.78%*** (0.68)	− 1.55%*** (0.21)	− 0.11% (0.67)
1995–2000	0.98% (0.84)	− 0.03% (0.36)	− 0.09% (0.64)	− 1.93%*** (0.51)
2000–2005	− 1.23% (1.02)	6.65%*** (1.49)	− 2.33%*** (0.35)	4.13%** (1.69)
2005–2010	− 1.25% (0.84)	− 0.33% (0.50)	− 1.86%*** (0.26)	− 1.76%*** (0.38)
Multiple-Use Protected Areas				
Overall park effect	− 0.31% (0.96)	2.39*** (0.72)	− 1.12% (1.26)	3.70%* (2.15)
1995–2000	− 3.32%*** (1.08)	0.60% (0.47)	− 2.98%*** (0.41)	1.27% (1.37)
2000–2005	− 0.03% (1.43)	5.68%*** (1.77)	− 1.03% (1.76)	6.71%** (3.18)
2005–2010	− 0.45% (0.50)	0.86% (0.76)	− 1.25% (1.11)	3.26% (2.01)

Note: Standard errors in parentheses. Time period effects (1990–1995, etc.) are estimated from equation [4]. The overall park effect is estimated after dropping the interaction term from equation [4] and rerunning the regression. Cluster robust standard errors are used in all regressions to control for spatial and serial correlation at the plot level. Since there is only one park in 1990–1995 (see Table 1), we do not report park effects in 1990–1995 when splitting the sample.
* $p<0.1$; ** $p<0.05$; *** $p<0.01$.

bance by about 2 percentage points compared to control observations in 2000–2005 and 2005–2010 and have a statistically significant overall park effect of about 2 percentage points. There are large qualitative and quantitative differences across fixed- and random-effects estimators in Table 5, and we come back to this issue in the discussion. For roads we did not find many differences across parks located closer to or farther away from nearest roads: both were negative and statistically significant in 2005–2010, with an impact of about 2 percentage points (results not reported in table).

Similar to strict protected areas, we find few statistically significant impacts of multiple-use protected areas on forest disturbance (Table 4). The overall effect from the fixed-effects estimator is positive but not significant. The estimated effect is insignificant across most time periods, with the exception of the 1990–1995 time period, where we find a positive effect of about 4 percentage points. There is only one park in this period (Table 1), so this effect is specific to only that park. Between the random- and fixed-effects estimators, the qualitative results are similar. When we stratify multiple-use parks by distance to Moscow, we find that overall, parks closer to Moscow (higher pressure) experience an increase in forest disturbance compared to areas outside of parks (Table 5); however, the only time period that is statistically significant is the 2000–2005 period. The effect of parks farther from Moscow (lower pressure) is mostly negative, although significantly different from zero only during the 1995–2000 time period. We do not find any difference across multiple-use parks located closer to or farther away from a road (results not reported in table). Similar to the estimates for the strict protected area sample, these stratified estimations do not include the one multiple-use park created in 1990–1995 (Table 1).

For both strict and multiple-use protected areas, we conducted a series of tests to verify the robustness of these results, including the visualization of parallel trends in Figure 2, placebo tests of the impact of protected area observations before protection occurred, and falsification tests in which we randomly assigned "treatment" to control observations. Placebo tests did not indicate that our findings were susceptible to an over- or underestimation of treatment effects (also known as Ashenfelter's dip). However, we did find a statistically significant and positive effect of the park created in 1990–1995 (for strict and multiple use) and the park created in 2005–2010 (for multiple use) in some pretreatment periods. One reason for this positive pretreatment impact could be if the government or private firms preemptively logged areas designated for protection. Since these types of differences violate the parallel trends assumption, we dropped these two parks that had a significant effect in pretreatment periods and reran the regressions: the results in Table 4 were not materially different, and therefore, we present results including all parks. Finally, falsification of controls did not result in any statistically significant effects of "treatment" on forest disturbance at the 5% level or lower. Results for these tests can be obtained by request from the authors.

VI. DISCUSSION

We set out to do two things in this paper: (1) provide the first quantitative assessment of the effectiveness of protected areas in post-Soviet European Russia at reducing forest disturbance, and (2) evaluate whether including plot fixed effects into panel estimators (rare in the environmental program evaluation literature) generates significantly different estimates from estimators without plot fixed effects. We address each of these in turn below.

Unlike most studies of protected area effectiveness in the tropics (e.g., Andam et al. 2008; Pfaff et al. 2009, 2013; Nelson and Chomitz 2011), we do not find that protected areas in European Russia consistently reduce forest disturbance pressure relative to matched controls. This may be due to the large temporal changes in governance and funding of protected areas, and the overall socioeconomic shocks occurring in post-Soviet Russia. The direction of the impact of parks in post-Soviet European Russia, that is, whether the park has a negative or positive effect on disturbance, also varies by the type of park—strict or multiple use —and the time period. While studies of protected area effectiveness in the tropics find that the magnitude of the effect of protection varies by park type, due to differences in their location and allowable uses, none report parks having higher rates of disturbance than areas outside of protection.

In European Russia, strict protected areas are located in more remote locations, have more funding and better enforcement, and do not permit logging, compared to multiple-use areas. When they have a statistically significant impact on forest disturbance, they tend to lead to a decrease in disturbance compared to similar observations located outside of protection. While their impact may appear small—ranging from 1 to 2 percentage points over a 5-year time period—it is comparable to findings of a global evaluation of protected area effectiveness (Joppa and Pfaff 2011) and is reflective of the overall low rates of forest disturbance in our study region, which ranges between 2% and 6% for a 5-year time period between 1990 and 2010. The time periods when strict protected areas are effective are the periods with the lowest overall disturbance rates in our study (i.e., 1995–2000 and 2005–2010), and in general, disturbance within protected areas follows similar temporal trends as disturbance outside of parks (Figure 2). This suggests that when there is a lot of pressure, strict protected areas are not able to block these threats.

Additionally, unlike studies in the tropics that find that parks located closer to threats have higher impact estimates because they have more threat to block (e.g., Pfaff et al. 2009, 2013), we do not find much difference in effectiveness across strict protected areas located closer to or farther from Moscow or major roads. The only exception to this is the positive and significant coefficient on parks closer to Moscow in 2000–2005, a time period associated with higher forest disturbance in our study area and corresponding to the end of the Asian financial crisis. The positive and significant effect of parks located closer to Moscow in 2000–2005 supports the conclusion above that when pressure is higher, strict protected areas are not effective. At best, designation of strict protected areas in post-Soviet Russia has resulted in a 1 to 2 percentage point reduction in forest disturbance compared to unprotected forest plots.

Multiple-use protected areas in post-Soviet European Russia do not fare much better, with very few statistically significant differences and, at times, higher forest disturbance than comparable control observations. The reason for this is not a lack of forest disturbance threat—these parks tend to be located in locations more susceptible to logging and have high rates of disturbance—but the inability to block disturbance threats. This finding is in contrast to results from tropical country studies that show multiple-use areas have larger impacts on lowering deforestation than strict protected areas because there is more threat to block (Pfaff et al. 2013), or such areas are at least as effective as strict protected areas (Nolte et al. 2013). The low effectiveness of Russia's multiple-use areas may be due to federally permitted logging leases to private firms or may be indicative of illegal activity, such as the sanitary logging practice conducted by the Federal Forest Service. As noted in Section II, both types of multiple-use areas permit logging, and there have been perverse incentives for local forestry officials to allow logging on federal lands to generate revenue for their own budgets. Of course, there is also a shortage of funding and management noted for multiple-use areas in post-Soviet Russia, indicating that illegal logging within these boundaries is possible. While it appears that parks located closer to Moscow are more likely to have forest disturbance within their borders, this does not shed much light on whether this reflects legally permitted (since transportation costs would be lower) or illegal (since pressure to take logs would be higher) disturbance. Again, this higher rate of forest disturbance closer to Moscow occurs in a time period (2000–2005) of increased disturbance rates across European Russia's forests.

Our estimated impacts of post-Soviet European Russia protected areas are important to bear in mind as a recent gap analysis for conservation in Russia has proposed the creation of an additional 403 federally protected areas (Krever, Stishov, and Onufrenya 2009). Multiple-use protected areas, such as national parks and federal zakazniks, make up the majority of the proposed new protected areas. Our results raise questions about enforcement against illegal logging and or possible permitted logging operations within multiple-use areas. At best, current multiple-use areas are a zero sum game in that they neither increase nor decrease forest disturbance relative to similar areas outside of protection. Before any new multiple-use areas are created, there is a need for on-the-ground research to understand why these park types appear to be susceptible to forest disturbance.

We turn next to the methodological evaluation of including plot fixed effects into panel estimators for impact evaluation. The number of impact evaluations is growing in the conservation and environment field (Pattanayak, Wunder, and Ferraro 2010; Ferraro 2011). The most common approach is to use matching to construct a valid control group and then cross-sectional regression to estimate the treatment effect. Some recent studies using this method to estimate the impact of protected areas include those by Andam et al. (2008), Pfaff et al. (2009, 2013), Ferraro and Hanauer (2011), Joppa and Pfaff (2011), Nelson and Chomitz (2011), and Nolte et al. (2013). There is also an increasing interest in using program evaluation methods to estimate the impact of payments for ecosystem services programs (e.g., Pfaff, Robalino, and Sanchez-Azofeifa 2008; Uchida, Rozelle and Xu 2009; Robalino and Pfaff 2013), and some of these studies have used difference-in-difference and fixed-effects methods to estimate treatment effects on forest protection (see Alix-Garcia, Shapiro, and Sims 2012; Arriagada et al. 2012). A relevant question is whether moving to the fixed-effects structure is critical for robust estimates of treatment effects. While random-effects estimates differ from cross-sectional regression in the modeling of serial correlation in the unobservables, the identification assumptions are identical across both estimators. We interpret random-effects estimates of protected area effectiveness as methodologically similar to the conventional environmental program evaluation literature.

The inclusion of plot fixed effects generates differences in qualitative and quantitative effectiveness estimates for protected areas, especially when the samples are stratified by distance to Moscow. Hausman tests confirm the statistical difference across fixed- and random-effects estimates (5% level).[5] While the

[5] We use cluster robust standard errors, so the conventional Hausman test of random versus fixed effects is incorrect since it relies on an assumption that random effects are

magnitude of the difference in the effectiveness estimates between fixed and random effects seems small in absolute probability terms, fixed-effects estimates are 1.4 times larger for 2005–2010 for strict protected areas and 1.2 times larger for 1990–1995 for multiple-use protected areas (Table 4). Taking the 2005–2010 strict protected area estimates and park size as an example for context, the effectiveness estimates from fixed effects indicate that strict protected areas prevent approximately 130 km^2 of disturbance over a 30-year time horizon, while the corresponding random-effects estimates indicate only 92 km^2. Since both estimators use the same sample, differences between fixed and random effects imply that time-invariant plot unobservables are correlated with protected area status. The presence of such unobservables can bias post-matching regression estimates of conservation effectiveness; whether this will result in an under- or overestimation of impacts depends on the unobservables.

While bias from time-invariant unobservables could be reduced in cross-sectional post-matching regressions by collecting more data on plot characteristics or instrumenting for protection, such data are not always easily available or well measured, and protection instruments are often far from obvious (see Sims [2010] for an example of an instrumental variables approach to protected area impacts). Our results suggest that building better temporal variation with spatial land use/land cover data can reduce the number of assumptions required for identification of protected area—or other environmental program—effectiveness. The identification of park effectiveness with plot fixed effects relies on (1) repeated remote-sensing landscape images over time, and (2) temporal variation in the location of protected areas (or other environmental program) within the time frame of the estimation sample. The incorporation of similar panel methods into evaluations of environmental policy on land use is becoming increasingly feasible given release of the Landsat archives (Goward et al. 2006; Blackman 2013) and the advancement of remote-sensing techniques to provide temporally rich land cover change classifications (Huang et al. 2008, 2009).

efficient. We conduct a robust Hausman test using the linear correlated random effects model with cluster robust standard errors (see Wooldridge 2010, 332), which is also known as the variable addition test for fixed versus random effects.

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