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ABSTRACT *Pollution from agricultural fires is a global health issue that is particularly challenging where smallholders depend on burnings for subsistence. In Acre state, Western Amazon, a partial ban on fire, enforced with fines, is coupled with subsidized tractors. To evaluate this policy, a discrete choice experiment and contingent valuation were merged into a novel statistical variant of the Hicks-Kaldor test that is robust to preference heterogeneity. Among 27 ways to extend the ban, 5 could improve both respiratory health and smallholders' welfare, when compensated with tractors that are available for longer hours and at the right time of the year.* (JEL Q51, Q52)

1. Introduction

The use of fire in agriculture is a global health issue. Every year around the world, crop residues and debris from slashed vegetation are burned to clear fields for agriculture. The chief benefit of burning is that it is inexpensive and can help with soil fertilization and weed management. However, agricultural fires release pollutants that have major consequences for respiratory health. A recent policy brief by the World Bank (Cassou 2018) identified China, India, the United States, and Brazil as the top four burners of agricultural residues worldwide. In China and India, pollutants from agricultural fires cause cough and eye irritation, exacerbate chronic respiratory illnesses, and reduce lung function (Chen et al. 2017; Shi et al. 2014; Kumar and Kumar 2010; Agarwal et al. 2012). In the United States, pollutants released by the burning of crop residues and rangeland may amount to 33% of annual county-level emissions of fine particulate matter (Pouliot et al. 2017, figure 2; McCarty 2011). The Southeast Asian haze has health impacts that reach disastrous proportions in El Niño years (Tacconi 2016; Ramdhan et al. 2018). This large-scale air pollution problem is caused by burnings conducted to clear native vegetation for tree plantations and also for subsistence agriculture (Tacconi and Vayda 2006; Marlier et al. 2015; Field et al. 2016).

In the Brazilian Amazon, landholders slash and burn forest, fallow vegetation, and also degraded pasture, seeking profit from cattle and cash crops but also food security (Tasker and Arima 2016; Nepstad, Alencar, and Moreira

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Appendix materials are freely available at http://le.uwpress.org and via the links in the electronic version of this article.

1999). As a consequence of agricultural burning, the region's inhabitants suffer from a wide range of respiratory-related problems (e.g., cough, chest pain, sore throat, fever), and the impacts of other chronic respiratory diseases and infections are exacerbated (e.g., asthma, bronchitis, pneumonia, chronic obstructive pulmonary disease); for more details, see Appendix A, Sections 1 and 6. Children and the elderly are the most affected by fire-related respiratory problems, with reduction in lung function detected in the former (Nunes, Ignotti, and Hacon 2013; Mascarenhas et al. 2008; Jacobson et al. 2012). It is estimated that the Brazilian mortality rate from pollution caused by agricultural fires is statistically equivalent to the mortality rate from diseases of the blood, including HIV (Reddington et al. 2015, table S3; SUS 2018).

A particularly worrying case in Brazil is that of Rio Branco, a western Amazon town with 330,000 inhabitants in the state of Acre. Every year the residents of Rio Branco are exposed to smoke, most of which is caused by fires used to prepare the land for agriculture (Silva et al. 2016; Do Carmo, Alves, and Hacon 2013; Duarte, Júnior, and Mesquita 2007). During the 2014 "fire season" (August to November), concentrations of fine particulate matter[1] within a 100 km radius of the Rio Branco town center exceeded the safety limit recommended by the World Health Organization more than 90% of the time. In 2015, the safety limit was exceeded more than 99% of the time. Fine particulate matter is one of the most health-damaging air pollutants (Arbex et al. 2014; Silva et al. 2016; Reddington et al. 2015). There is now well-documented epidemiological evidence of major health impacts in Rio Branco and surrounding areas from agricultural fires (Silva et al. 2016; Mascarenhas et al. 2008; Do Carmo, Alves, and Hacon 2013; see also Appendix A, Sections 1 and 6).

Many countries have responded to the severe health consequences of agricultural fires by banning the practice. This is true for most of Europe, India, and China (Cassou 2018; Tallis et al. 2017; Xia 2014; Pan, Crowley, and Lehmann 2011). However, in developing countries, compliance with bans has been low and enforcement has proved very difficult (see Theesfeld and Jelinek [2017] for Russia; Pan, Crowley, and Lehmann [2011] and Xia [2014] for China; and Tallis et al. 2017 for India). For instance, in Xinyi, northern China, the local government deployed 500 policemen to the fields to prevent burnings. This strategy proved ineffective, however, as farmers responded by burning in the evenings, at the end of the police work day (Xia 2014). In addition to the high monitoring costs associated with enforcing fire bans in developing countries (Keck and Hung 2018), one of the main factors preventing bans from being fully effective and socially acceptable is the relevant role played by poor smallholders. This group, which conducts a substantial proportion of the burnings, has limited access to labor, fertilizers, machinery, and finance and relies on fire for subsistence (Tacconi 2016; Carmenta, Coudel, and Steward 2018; Keck and Hung 2018).

The fire dependence of smallholders has motivated a second generation of policies in developing countries to couple (partial) bans with subsidies for alternatives to fire. In northern India, where 23 million tonnes of rice residues are burned every year (Tallis et al. 2017), the current policy bans burning and subsidizes a planting procedure that mixes mechanization and residue mulching, known as "Happy Seeder" (Tallis et al. 2017). In southern China, the local government of Xinyi has subsidized a machine to shred crop residues to avoid peasants burning such by-products (Xia et al. 2014). Additionally, in Punjab (northern India), no-till soil conservation has been subsidized as an alternative to fire (Kumar, Kumar, and Joshi 2015; Cassou 2018).

To summarize, understanding how to balance incentives with command-and-control measures to reduce respiratory damage caused by agricultural fires without threatening the well-being of poor, fire-dependent smallholders is a policy-relevant issue. To address it, we collected data on preferences for fire policy and health protection from rural and urban residents of Rio Branco and surrounding areas.

[1] PM 2.5 concentrations were estimated by the CATT-BRAMS atmospheric model (Freitas et al. 2009) and were supplied by Demerval Soares Moreira (FC/UNESP) and CPTEC-INPE, the Brazilian atmospheric science institute. The safety limit of World Health Organization is 25 μg/m3.

This region provides an internationally outstanding case study not only due to high levels of fire-induced pollution, but also because it has pioneered a policy approach combining a partial fire ban with subsidized tractors, which smallholders can use to clear fields instead of burning (details in section 2.1).

We surveyed rural fire users and urban victims of fire pollution using stated preference methods. Adult urban residents completed a contingent valuation survey that asked their willingness to pay (WTP) to avoid respiratory illnesses affecting them or their children. Rural (smallholder) fire users completed a discrete choice experiment that elicited their willingness to accept (WTA) policy amendments that would reduce their right to burn but increase their access to tractors. We undertook two main steps to aggregate individual values resulting from contingent valuation and the discrete choice experiment. First, we estimated the marginal increase in respiratory illnesses caused by fires using econometrics and GIS techniques. Second, we devised a statistical version of the Hicks-Kaldor test to mitigate uncertainty due to preference heterogeneity and then assessed the efficiency of policy amendments.

2. Rationale for Cost-Benefit Analysis

Study Region and Policy Context

The study region (SR) is the geographical area within 100 km of Rio Branco, the capital municipality of Acre state in the Brazilian Amazon (Appendix B, Figure B1; 25,300 km^2, approximately the size of Macedonia, population > 400,000 [IBGE 2013]). The 100 km threshold was established based on the budget available for field surveys and human-supervised mapping of burn scars. Notwithstanding, the census tracts within the SR account for 58% of the state's population (IBGE 2013).[2] The SR was subdivided into nine "sectors" to capture heterogeneity in fire use intensity and in the effects of fire on health. Divisions were made on the distance and direction to the urban town of Rio Branco (Figure B1 in Appendix B and Appendix A, Section 3.1.1) and with a spatial resolution of 10 km (in order to maintain computational tractability).

The Acre state government has made several attempts to control agricultural fires since 1999 (Brown et al. 2011; CEGdRA 2013; Acre 2013). These attempts have included improved remote monitoring and early warning systems, and the expansion and optimization of fire brigades. Policy instruments targeted at changing the behavior of fire users, including direct regulation and economic incentives, have also been implemented. In what follows, we provide an overview of these policy instruments as applied in Acre state recently.

In 2009, an attempt to control agricultural fires through direct regulation was made. This took the form of a ban on fire use—announced by federal and Acre state public prosecutors—that was to take effect in 2012. However, the Acre state government and its environmental institute withdrew the ban in 2012. The setback was due to the social cost the ban imposed, as smallholders and indigenous people lacked the means to secure their subsistence without relying on fire to grow crops (Acre 2013). Following this ban, a less restrictive policy was adopted and is still in place. It allows fire use only by smallholders and indigenous people, although it restricts the area that can be burned to 1 ha/landholding/year. Under this policy, larger areas cannot be burned, and transgressions incur heavy fines—ranging from \$153/ha (10% of local average crop production value per hectare) to \$1,534/ha (100% of production value)[3] (IBGE-PAM 2018). In addition, smallholders must obtain formal permission to burn from the government.

Programs based on multiple incentives have also been undertaken in Acre state to control deforestation and to reduce fire use by smallholders (see Appendix B, Table B1). These programs have included government support for rural extension and technical as-

[2] It should be stressed that the two subpopulations considered, smallholders and residents of the urban center of Rio Branco, corresponded to 70% of the SR's population of fire users and 73% of the SR's total population, respectively.

[3] Personal communication with the head of the licensing division of Acre state's Environmental Institute (IMAC), January 2017.

sistance, mechanized land preparation, and, in some cases, input subsidies. Conditional payments were also attempted. These payments required progress to be made to decouple land management from deforestation and fire use, but (ex post) were found to be incompatible with the available budget.[4] Currently, two main incentive instruments are in use in Acre state, both based on in-kind subsidies (Table B1).[5] The first instrument promotes fire reduction through land use change, by targeting fish farming (e.g., via fish tank building) and agroforestry (e.g., via input donations) as alternatives to the staple crops traditionally associated with fire. The second instrument encourages land management change by offering subsidized tractors that are mainly used in the cultivation of staple crops. According to field observations, the second instrument is currently supporting the largest number of smallholders to abandon fire.

Tractor subsidization was initially introduced to support implementation of the existing partial ban on burnings. It also allows smallholders (with landholdings of 47 ha on average) to cultivate areas larger than 1 ha not only for subsistence but also for income.[6] Both Acre state and some of its municipalities, including Rio Branco, have their own tractor fleets. These fleets provide a mechanized land preparation service for smallholders, generally in two stages.[7] In the first stage, tree stumps, which are widespread, are removed with a bulldozer. This makes the ground accessible for tractors on tires, which, in the second stage, harrow the soil. This land preparation service is supplied by government-employed technicians, who schedule the tractor supply and drive the machines. In this case, the government maintains the tractors and smallholders have to provide only fuel and food for the drivers. Alternatively, tractor ownership can be temporarily transferred to smallholder communities, which become responsible for their maintenance. Roughly 850 to 950 households of smallholders per year benefited from this scheme between 2015 and 2016, with nearly 800 to 980 ha per annum being removed from burning.[8]

In summary, when applied separately, neither direct regulation in the form of an unconditional ban nor conditional cash incentives proved feasible or effective at reducing fire use in Acre state. However, a combination of "milder" versions of these two policy instruments—a partial ban coupled with an in-kind agricultural subsidy in the form of tractor-usage—seems to have been relatively successful. This "sticks-and-carrots" approach could be intensified by coupling an increase in tractor supply with a decrease in the legal burning limit. This change could potentially generate benefits associated with (1) avoided respiratory illnesses, and (2) expanded access to tractors, which would compensate smallholders for the costs imposed by the reduced right to burn.

To test whether such a favorable outcome is realistic, a cost-benefit approach (CBA) based on stated preferences was developed. The stated preference technique is appropriate for measuring not only health benefits (as revealed by a wide literature, e.g., Alberini et al. 1997; Richardson, Loomis, and Champ 2013; Ara and Tekeşin 2016), but also the benefits and costs provided by subsidized tractors and constraints on burning. Smallholders' decision to replace fire usage with tractors is driven not only by objective profit levels, but also by a number of factors related to preferences. These include risk aversion and subjective discount rates (Bowman, Amacher, and Merry 2008; Mburu et al. 2007; Angelsen 1994). In addition, learning costs and disutility associated with the labor effort required by

[4] Personal communication with representatives of the Acre state agricultural support department (Seaprof), and conversations with the heads of the institutions participating in the farm certification and sustainable development programs, January 2017.

[5] Information gathered during visit to Safra (Municipal Department of Agriculture, Subdivision of Production, Tractor Supply Program, Rio Branco municipality), December 2016. Personal communication with representatives of the Acre state agricultural support department (Seaprof), and conversations with the heads of the institutions participating in the farm certification and sustainable development programs, January 2017.

[6] Information gathered during visit to Safra, December 2016.

[7] Information gathered during visit to Safra, December 2016.

[8] Information gathered during visit to Safra, December 2016.

fire-based and fire-free land management also play a role (Schuck, Nganje, and Yantio 2002; Carmenta, Coudel, and Steward 2018; Mendoza-Escalante, Börner, and Hedden-Dunkhorst 2003). Finally, there may also be deep-rooted beliefs regarding the pros and cons of fire-based agriculture (Theesfeld and Jelinek 2017; Carmenta et al. 2013). However, basing a CBA upon stated preference techniques also presents methodological challenges, as applied studies suggest that preferences are significantly heterogeneous (see below). This heterogeneity can lead to large variance in costs and benefits, which could be significant enough to considerably reduce the accuracy of a standard CBA.

The Conceptual Foundation of the CBA

To understand how the applied literature has dealt with the challenge of preference heterogeneity, we reviewed eight applied studies and two CBA textbooks. This literature covered the topics of preference heterogeneity and extrapolation of individual values (i.e., benefit transfer and aggregation). Three studies (Colombo, Calatrava-Requena, and Hanley 2007; Amador, González, and Ortúzar 2005; and Brouwer, Martin-Ortega, and Berbel 2010) captured preference heterogeneity at the most fine grained level possible: that of individuals. The third study strongly supports the individual-level perspective, showing that failure to account for preference heterogeneity at the individual level underestimated WTP for water quality by 33%. Borghi (2008) introduced heterogeneity in CBA by reporting results that excluded or included nonusers of the valued good. Direct and passive user groups were also the source of preference heterogeneity in the CBA of cultural heritage by Báez and Herrero (2012). Heterogeneity was captured at the level of social groups by the CBAs of Pearce, Atkinson, and Mourato (2006, chap. 15) and Brent (2007), both of whom assumed that preferences were potentially different between low- and medium- to high-income individuals. Three studies incorporated spatial-level heterogeneity in their preference function. Schaafsma et al. (2013) captured spatial heterogeneity in preferences for water-related recreation based on locational variables such as feasible recreation activities, population-site distance and direction, and local substitute availability. In a similar fashion, Bateman et al. (2006) incorporated both locational and socioeconomic heterogeneity within a WTP function. Finally, the CBA of Hyytiäinen et al. (2015) also captured heterogeneity at the geographical level (among countries in the Baltic Sea), as well as in terms of self-selected goods valued. Pearce, Atkinson, and Mourato (2006) were the only ones who explicitly mentioned a procedure to incorporate variance of preferences in CBA.

In summary, most of the reviewed studies captured preference heterogeneity between social groups or geographical domains but missed potentially relevant heterogeneity within such aggregates at the level of individuals. This contrasts with results from "state-of-the-art" modeling of discrete choice (Hole and Kolstad 2012). These results reveal significant variance, at the individual level, in estimates of preference for environmental and policy attributes (Kuhfuss et al. 2015; Christensen et al. 2011; Villanueva et al. 2016; among many others). If this variance is not pure noise, but instead reflects how individuals react differently to changes in the provision of the valued good, extrapolation based exclusively on the means of individual values elicited by stated preference surveys will be subject to error. Indeed, the Hicks-Kaldor test that provides the grounding for most applied CBAs is based on a product of averages of individual values by population size (Bockstael and Freeman 2005; Pearce, Atkinson, and Mourato 2006; Bateman et al. 2006; Borghi 2008), as follows:

$$\textit{Welfare surplus} = \hat{\Omega}_{\mathrm{B}} \frac{1}{N_{\mathrm{B}}} \sum_{i=1}^{N_{\mathrm{B}}} B_i - \hat{\Omega}_{\mathrm{C}} \frac{1}{N_{\mathrm{C}}} \sum_{j=1}^{N_{\mathrm{C}}} C_j, \quad [1]$$

where $\hat{\Omega}$ represents the size of better-off (B) and worse-off (C) populations, and B_i and C_j are specific WTP and WTA values, respectively. The estimator in equation [1], which is consistent in the statistical sense (Appendix B), nevertheless has a nonnegligible probability of producing point estimates significantly different from the "true" value of the target parameter. This error is more likely to occur the higher the degree of preference heteroge-

neity (according to Chebyshev's inequality; Greene 2003, theorem D.2).

Consequently, when a CBA is based on heterogeneous stated preferences from surveys with small samples, it must explicitly address sampling error to avoid incorrect policy prescriptions.[9] To do that, the CBA applied in this paper is based on a novel statistical version of the Hicks-Kaldor test. This approach proposes that population surplus is positive only if the point estimate exceeds a threshold, which is directly proportional to potential sampling error and, thus, to the estimator's variance. We calculate this variance in two steps: first, by bootstrapping better- and worse-off samples to estimate average benefits and costs, and second, by creating all combinations between estimates from the two samples to generate a large set of point estimates for the surplus (details are provided in Section 3).

In this paper, we evaluate potential policy amendments to Acre's fire-for-tractors policy. The specific amendments we analyze are quantitatively different ways of combining a reduction in the legal burning limit with an expansion in the supply of subsidized tractors. These amendments will impose costs on rural smallholders that depend on fire use, and yield benefits to urban dwellers that are affected by atmospheric fire pollution. The estimate for the size of the "worse-off" population, $\hat{\Omega}_C$, was the number of smallholders who burn areas above the legal limits established by the policy amendments (half and zero hectare). The size of the "better-off" population, $\hat{\Omega}_B$, was estimated as the number of respiratory illnesses avoided by policy amendments. Section 3 describes how the two estimates were obtained. Further details are provided in Appendix A, Sections 2 and 3.

[9] This phrase is rigorously correct as it refers to the case in which CBA is based on small and fixed samples of stated preference surveys. However, for sample size tending to infinity or approaching a satisfactorily large level, sampling error converges to zero and can be thus disregarded (in line with the consistency result of Appendix B). This second possibility is related to the Arrow and Lind (1970) theorem, which proposes that, given certain assumptions, public investment can be optimally decided solely on the basis of expected return without concern for variance. The theorem is thus not being disputed by the statement in the main text; we thank one of the reviewers for mentioning the relationship between our statement and the Arrow-Lind theorem.

3. Method and Data

Benefit of Fire Reduction

The expected degree of unfamiliarity bias (in the sense of Seip and Strand [1992]) introduced by a predefined good was higher for the better-off population due to the large diversity of respiratory illnesses and of the symptoms they cause. To minimize this bias, the object of valuation was defined as respiratory symptoms that urban dwellers had personally suffered recently (following Alberini et al. [1997] and in accordance with Schläpfer and Fischhoff [2012] and Boyle, Welsh, and Bishop [1993]). The personalized nature of the good valued was not feasible in a discrete choice experiment framework where the object of valuation is predefined by the analyst. The personalized good was controlled for in the WTP function according to measures of experienced symptoms and pain. For details on sampling and the questionnaire, the reader is referred to Appendix A, Section 4.

Concept of the Survey

The objective of the contingent valuation survey was to elicit the willingness of residents of the urban town of Rio Branco to pay to avoid recurrence of their most recent episode of respiratory illness. A total of 488 valid interviews were conducted. These interviews were guided by a questionnaire and began by gathering data on the reference illness, namely, the interviewee's most recent respiratory illness episode, and then proceeding to a dichotomous choice WTP question (Appendix A, Section 4.3). Two versions of the questionnaire were applied, one for adults and the other for children and teenagers. In the first version, the WTP question was addressed to the patient him/herself, and, in the second, to the individual financially supporting the infant patient; according to Alberini et al. (2010), this is the least biased way to elicit preferences for children's health.

A private good was used as the payment vehicle: the purchase of a vaccine capable

of providing full immunization against the reference illness. This is in line with previous studies that offered a vaccine as a means to elicit preference for health benefits and that achieved theoretically valid results (see, e.g., Morris and Hammit 2001; Araña and León 2002; Hadisoemarto and Castro 2013). High uptake during influenza vaccination campaigns (over 80% for nearly all targeted groups) attests to the credibility of this payment vehicle in the study region. Also, campaigns are often time limited (less than two months) or targeted to high-risk groups (≤35% of population), signaling that vaccination was costly and adding to the evidence that paying for the vaccine was also credible. The payment vehicle was well received by local health professionals, and only a small share of respondents reacted negatively to it (4%, see Section 4).

Econometric Approach

The theoretical validity of the survey was evaluated based on both single- and double-bounded dichotomous choice formats. However, the second is not reported or used for the CBA due to its inconsistency with rational preferences and associated downward bias (Bateman et al. 2001; Cooper, Hanemann, and Signorello 2002; Carson and Groves 2007).

The covariates of the WTP function were selected according to common practice when using contingent valuation to investigate respiratory illnesses (Alberini et al. 1997; Ortiz et al. 2011; Ara and Tekeşin 2016; Richardson, Loomis, and Champ 2013) and included illness duration, number of symptoms and pain experienced, age, household income, and additional controls (see Table 2 in Section 4). Three econometric models were estimated. First, a linear WTP function with disturbances following a Gaussian distribution with zero mean and constant variance was estimated with the maximum-likelihood command, as by López-Feldman (2012).[10] The second and third econometric models adopted nonlinear WTP functions specified according to assumptions regarding the distribution of disturbances. Both logistic and standard Gaussian distributions were considered, yielding, respectively, logit and probit models (see Carson and Hanemann 2005, 853; Richardson, Loomis, and Champ 2013). These two last models were run using code developed by Nakatani, Aizaki, and Sato (2016)[11] and also differ from the linear WTP function because they include bids as covariates.

Cost of Fire Reduction

Under the policy change evaluated in this research, the worse-off population (rural smallholders) would face reductions in the amount of land that is allowed to be burned and increases in tractor availability. Smallholders are very familiar with both of these land management tools, which are widespread in the SR. Thus a conventional discrete choice experiment embodying predefined goods was a reasonable approach to adopt. A discrete choice experiment also allowed respondents to apprehend, in a straightforward way, the multiattribute nature of the fire-for-tractors policy and allowed us to generate multiple combinations of attribute levels to evaluate as potential policy amendments. Details on the sampling procedure and questionnaire are found in Appendix A, Section 5.

Concept of the Survey and Attributes

The objective of the discrete choice survey administered to smallholders was to elicit their WTA compensation for a reduction of the area they are allowed to burn annually in exchange for expanded tractor support. In each choice set, respondents were asked to choose between the status quo (no change to current policy) and two alternative policy scenarios. The policy scenarios consisted of concurrent fire-for-tractors changes. First, the cap on annual burning limits would be reduced below the status quo level. Second, the subsidized tractor hours would be expanded above the status quo level. Additionally, a five-year annual payment was offered as compensation

[10] In fact, for the single-bounded model, the "singleb" routine was used, which is also mentioned in the cited paper.

[11] Logistic and log-logistic models returned exactly the same estimates. All other available specifications for disturbances' distribution (normal, log-normal, and Weibull) did not converge.

Table 1
Attributes and Levels for Each Option

Attribute	Levels
Allowed annual burning limit (ha/smallholding/year)	1,[a] 0.5, 0
Increase in bulldozer tractor (hours/year)	No increase,[b] +3 hours, +6 hours
Increase in harrowing tractor (hours/year)	No increase,[b] +2 hours, +8 hours
Delay in tractor arrival[c]	Current level of delay (=1),[b] no delay (=0)
Compensation payment (R$/smallholding/year)	R$ 0,[a] R$ 500, R$ 1,000, R$ 4,000, R$ 6,000

[a] Status quo–only level.
[b] These levels occurred in status quo but also in non–status quo options.
[c] This attribute was phrased in the survey as "reduction in tractor arrival delay."

for a reduction in the burning limit. The discrete choice experiment attributes are listed in Table 1, and an example of a choice card is found in Appendix A, Figure A5.

The "delay in tractor arrival" attribute was suggested by participants of the pilot surveys. Including this attribute added to the realism of the survey, as delays are often experienced under the current policy, and many respondents complained about the delays. A binary format for this attribute proved easier to propose to respondents.

Experiment Design and Econometric Approach

A fractional factorial D-optimal design was generated with the command by Hole (2016), including a binary for the status quo alternative (in line with Kuhfuss et al. 2015; Villanueva et al. 2016; Christensen et al. 2011). From this design, three different blocks of choice sets were obtained, each with six choice sets of three alternatives. This procedure was adopted both in the second pilot and in the final survey. In both surveys, each block—and consequently each choice set—was repeated more than six times, as recommended by Bunch and Batsell (1989, cited by Louviere, Hensher, and Swait 2000, 103). There were no dominant alternatives in the choice sets.

Following common practice (Hole and Kolstad 2012; Greene and Henscher 2010; Kuhfuss et al. 2015; Christensen et al. 2011; Villanueva et al. 2016) we estimated the following econometric models: conditional logit (CLogit), mixed logit in preference-space (Mixed), mixed logit in WTP-space (Mix-WTP), and the generalized multinomial logit (GMNL). The Mixed and GMNL models allow for preference heterogeneity with random coefficients for attributes (Hensher, Rose, and Greene 2005). This feature allows welfare estimates to capture idiosyncratic responses to policy attributes.

Welfare Analysis

Equation [2] describes WTA amendments to the fire-for-tractors policy, and is consistent with the welfare underpinnings of the random utility model and commonly adopted simplifying hypotheses (Bockstael and McConnell 2007).

$$WTA = [\gamma_0(L_0 - L_1) + \gamma_1(bull_0 - bull_1) + \gamma_2(harr_0 - harr_1) + \gamma_3(del_0 - del_1) + \delta + \mathbf{x}K]/\beta, \quad [2]$$

where $L \equiv$ burning limit in hectares per year, $bull \equiv$ hours of bulldozer tractor supplied, $harr \equiv$ hours of harrowing tractor supplied, $del \equiv$ delay on tractor arrival (binary; see Table 1), and $\mathbf{x} \equiv$ a vector of respondent-specific covariates. Status quo and policy amendment levels of attributes are subscripted, respectively, with 0 and 1. Attributes' coefficients are γ_0 to γ_3 and δ for the status quo, K for respondent-specific, and β for the compensation payment. The levels of the mechanization support attributes were presented to respondents as deviations from status quo levels, denoted by $bull_1'$, $harr_1'$, and del_1'. Then,

$$WTA = [\gamma_0 (L_0 - L_1) - \gamma_1(bull_1') - \gamma_2(harr_1') - \gamma_3(del_1') + \delta + \mathbf{x}K]/\beta. \quad [2']$$

An undesired property of equation [2′] is that the perpetuation of delayed tractor arrival does not increase WTA, whereas the elimination

of delay decreases it, as γ_3 is expected to be negative (which was indeed the case). More precisely, it is assumed here that smallholders will comply with the lower burning limit even with a delay in tractor arrival. However, in reality, this is possible only if smallholders incur additional costs, for example, delaying farming or investing in other fire substitutes (e.g., private tractor fleets). Therefore, the current assumptions create a hidden cost of delay perpetuation. An ad hoc way[12] to account for this cost is to normalize the disutility caused by status quo delay to zero, that is, to force $\gamma_3 del_0 \equiv 0$ (only end state delay matters). This ensures that WTA refers to the monetary compensation of all losses, which strengthens the conservativeness of the CBA.

Due to the random nature of coefficients in the Mixed specification (and GMNL), they were not monetized through division by the monetary attribute's coefficient, but directly obtained from the MixWTP model (Hole and Kolstad 2012).

Test of Policy Efficiency

To test the null hypothesis that policy amendments are inefficient, a statistical version of the original Hicks-Kaldor test is proposed here (hereafter referred to as the SHK test). The unique addition to the original test is the idea that it can be understood as a test for the population or "true" value of surplus. The hypotheses in question are the following:

$$H_0: \mu_S = \Omega_B \mu_B - \Omega_C \mu_C = 0 \text{ versus}$$
$$H_1: \mu_S = \Omega_B \mu_B - \Omega_C \mu_C > 0, \qquad [3]$$

with population means and sizes denoted by, respectively, μ and Ω. According to the analogy principle, a statistic for this test is

$$\bar{s} \equiv \hat{\Omega}_B \frac{1}{N_B} \sum_{i=1}^{N_B} B_i - \hat{\Omega}_C \frac{1}{N_C} \sum_{j=1}^{N_C} C_j = \hat{\Omega}_B \bar{B} - \hat{\Omega}_C \bar{C}.$$

This represents the surplus estimated with sample data. The consistency of this estimator is discussed in Appendix B.

For the current hypotheses test, the probability of type I error is $P(\bar{s} > s_{0.05} | \mu_S = 0) = 0.05$, with $s_{0.05}$ being the critical value with 5% probability at its right. Therefore, if $\bar{s}_{OBS} > s_{0.05}$, the null is rejected, with $\bar{s}_{OBS}$ being the point estimate for surplus. However, $s_{0.05}$ cannot be directly calculated as the true distribution of the statistic $\bar{s}$ is unknown.[13] In fact, only one value of $\bar{s}$ is observed, as the available data consist of one benefit (WTP) sample and one cost (WTA) sample. Notwithstanding, the empirical distribution of the statistic can be obtained using the mixed bootstrapping-convolution approach of Poe, Giraud, and Loomis (2005) and the general bootstrap-based hypotheses testing approach of Efron and Tibshirani (1993, chap. 16). The implementation algorithm is as follows.

1. Generate the empirical distributions of $\bar{B}$ and $\bar{C}$, with 1,000 nonparametric bootstrap replicates in each case. Create all 1,000,000 pairs of $\bar{B}$ and $\bar{C}$ values (a full-combinatorial approach, in line with Poe Giraud, and Loomis 2005).
2. Generate the empirical distribution of $\bar{s}$ computing, for each pair, $\bar{s}^* = \hat{\Omega}_B \bar{B} - \hat{\Omega}_C \bar{C}$.
3. Take as a critical value the 95th percentile of the empirical distribution of $\bar{s}$ subtracted from $(1/M)\sum_{m=1}^{M} \bar{s}_m^*$ (M = 1,000,000). The subtraction imposes the null of H_0: $\mu_S = 0$. This step is a coarse adaptation of the way Efron and Tibshirani (1993, section 16.4) recommend the null to be imposed. The coarseness is dictated by the impossibility of directly bootstrapping values from the (unknown) distribution of $\bar{s}$. The critical value will be denoted as s_γ^*.
4. Calculate the observed value of the statistic with sample data, $\bar{s}_{OBS}$.

5a. If $\bar{s}_{OBS} \leq s_\gamma^*$, the null hypothesis of $\mu_S = 0$ is not rejected (inefficiency).

5b. If $\bar{s}_{OBS} > s_\gamma^*$, the null hypothesis is rejected and the alternative hypothesis of $\mu_S > 0$ is kept (efficiency).

[12] The non–ad hoc approach of including attribute interactions, capturing increase in delayed hours or reduction in land area prepared on time, failed by leading to incoherent WTA estimates due nonsignificance of interactions or of the burning limit attribute.

[13] Rigorously, the distribution of the statistic is given by the following convolution, which cannot be analytically derived: $f_{\bar{s}}(\bar{s}) = \int_{-\infty}^{\infty} f_{\tilde{B},\tilde{C}}(\hat{\Omega}_B \bar{B}, \hat{\Omega}_B \bar{B} - \bar{s}) d\bar{B}$, with $\tilde{B} \equiv \hat{\Omega}_B \bar{B}$ and $\tilde{C} \equiv \hat{\Omega}_C \bar{C}$.

This procedure was repeated for the 27 distinct non–status quo alternatives of the smallholders' discrete choice experiment, which are here understood as the policy amendments to be evaluated by the CBA. Amendments differ in the empirical distribution and in the point value of the statistic, but do not differ on $\overline{B}$ (as the preferences of the better-off population could be measured only in regards to their own health and not to a particular policy amendment).

Estimates of Population Sizes

Cost and benefit were aggregated based on two physical factors respectively: the total number of smallholdings above the burning limit, $\hat{\Omega}_c(p,s)$; and the number of respiratory illnesses avoided, given by the triple product $\hat{\Omega}_B(p,s) = C_a(s) \cdot C_b \cdot \Delta HS(p, s)$. Here, the letter p represents a specific policy amendment, which sets the burning limit level, while s indexes the nine sectors of the SR, defined in order to capture spatial variation in the impact of fire on pollution.

Better-Off Population

The impact of agricultural fires on respiratory illness level—attested to by epidemiological evidence (Appendix A, Section 1)—can be understood as occurring through two stages. In the first stage, fires impact the atmospheric level of pollutants. In the second stage, pollutants impact human health. An econometric approach was adopted to measure the intensity of impacts at each stage, following a suggestion by the British Treasury manual (see HM Treasury 2003, box 2.1) and multiple previous studies (Deryugina et al. 2016; Chagas, Azzoni, and Almeida 2016; Rangel and Vogl 2016; de Mendonça et al. 2004; Nicolella and Belluzzo 2015).

For the first stage, the following time series model was estimated for each sth sector of the SR:

$$AOD_t = \beta_0 + \beta_1 fire_{t,s} + \beta_2 d\_wind_{t,s} + \beta_3 fire_{t,s} d\_wind_{t,s} + u_{t,s},\ t = Jan2010, \ldots, Dec2016, s = 1, \ldots, 9, \quad [4]$$

with $AOD \equiv$ urban level of aerosol optical depth, $fire \equiv$ fire point detections, $d\_wind \equiv$ binary for wind blowing from sth sector toward the SR's urban area, and $u \equiv$ disturbance term. *AOD* is derived from smoke, which is a proxy for pollution (Mishra et al. 2015; Bevan et al. 2009). Data on smoke were sourced from NASA's MAIAC high-resolution aerosol product, in 1 km pixels (details given by Martins et al. [2017]). Data from fire point detections by NASA MODIS came from INPE (2017), in 1 km pixels, and data regarding wind from MERRA (2017), in 50 km × 65 km pixels.

The estimated coefficients were used to calculate the marginal effect of fires detected in each SR sector, most of them rural areas, on the AOD level of the urban town of Rio Branco—the latter belonging to the "urban area" sector. By statistically measuring the relationship between fires at rural locations and pollution in an urban location, many of the underlying forces are automatically captured. For example, synchronicity of peaks and valleys between fire (rural) and smoke (urban) time series tends to increase as (1) rural-urban distance decreases and (2) wind and topography become increasingly favorable to the transport of particles (Mishra et al. 2015).

For the second stage, the following panel dose-response model was estimated:

$$Y_{it} = \exp(\beta_0 + \alpha AOD_{it}' + X_i'\beta + d_{it}'\gamma + a_{it}) + u_t,\ i = 1, \ldots, 128,\ t = 1, \ldots, 12, \quad [5]$$

with i indexing urban area neighborhoods, t indexing fire season months (August to November of 2014 to 2016), Y being a count of respiratory illnesses, and $\mathbf{X}$ a vector of time-invariant covariates (detailed below). Binary variables for year and month are subsumed to the vector d_{it}.

To avoid counting illnesses caused by other factors, only fire season months are accounted for (as done by Silva et al. [2016] and Reddington et al. [2015]), as distinct sources of pollution release distinct amounts of specific pollutants, thus generating distinct health impacts (see Arbex et al. 2014).

Given that the dependent variable is a count, the functional form above corresponds to a Poisson specification. A negative binomial specification was also estimated, and for both specifications, the population mean, fixed effect, and random effect estimators

were applied. Nonspherical disturbances were treated by bootstrapping and clustering at the cross-section level (Cameron and Trivedi 2009, chap. 18; Wooldridge 2002, chap. 19).

The dependent variable, Y, was obtained from outpatient records of Rio Branco urban town. Eight definitions for Y were considered, including the total count, four age group counts (0–4 years, 5–17, 18–65, over 65), and three illness group counts (influenza, pneumonia, other illnesses). Therefore, combining all econometric specifications and dependent variables, a total of 48 models were fitted to the data, each bootstrapped 25 times; marginal effects were averaged across bootstrap repetitions.

The time-invariant covariates were selected from previous dose-response studies of air pollution (Deryugina et al. 2016; Chagas, Azzoni, and Almeida 2016; de Mendonça et al. 2004). They captured the availability of public health services (hospitals and outpatient units), a proxy for vehicular sources of pollution (density of streets), population (density and age groups), income, and a proxy for wealth (density of nightlights; see Henderson, Storeygard, and Weil 2012).

Finally, the last component needed to aggregate benefit, $\Delta HS(p,s)$, required estimates of the number of *point* fire detections avoided by policy (in consistency with the first stage detailed above). However, the aggregation of cost required estimates of burned *area* (as the burning limit is expressed in hectares). To reconcile these two different measures of fires, and ensure proportionality at the scale of the SR sectors, an econometric model was employed. It explained fire point detections as a function of burned area, an intercept, and year dummies, at smallholding level and with annual panel data from 2014 to 2016. This was used to predict fire point detections for two scenarios: (1) under the status quo, with observed burned areas kept unchanged, and (2) under policy amendment, with burned areas of all smallholdings being reduced to the amended burning limit (half or zero hectare per smallholding). The difference between predictions for the two scenarios was the estimate for $\Delta HS(p, s)$ (more details in Appendix A, Section 3.1.3).

Worse-Off Population

The worse-off population was estimated as the number of smallholdings with burned area above the new limit to be introduced by the pth policy amendment. This was a GIS-based count built on (1) burned-area maps made with supervised classification of Landsat-8 (30 m resolution) and Resourcesat (56 m) images for the peak months of the 2014 to 2016 fire seasons (details by Silva 2017, chap. 1), and (2) polygons demarcating smallholdings (≤400 ha) from the database of the Brazilian Rural-Environmental Registry (SICAR 2017).

4. Results

Benefit of Fire Reduction

Description of Survey Data

The sample was representative of the Rio Branco municipal population, as detailed in Appendix A, Section 4.2. The average reference illness lasted for nine days and caused 10 symptoms (Table 2). Most respondents (75%) reported a moderate level of pain and discomfort and acquired the reference illness outside of the fire season, most commonly in March to May (51%). Only 36% of respondents went to a doctor and received a diagnosis. Among them, the most recurrent respiratory disease was influenza. All symptoms with high frequency (>50%) affected the nose, caused headache, sore throat, cough (dry and with phlegm), fever, and malaise. The symptoms reported by respondents were consistent with previously documented reports in the literature describing the effect of fire on respiratory health and also with outpatient data that were the basis for the aggregation of WTP (details in Appendix A, Section 6).

There were 23 protest responses to the WTP question, and 31 observations were additionally discarded due to inconsistent data, interview failures, or preference inconsistency (in the latter case, open-ended WTP was less than the maximum accepted bid). A total of 488 observations—268 adults and 220 children—were considered for econometric analysis.

The payment vehicle was credible for most respondents, as attested by the 4% protest rate,

Table 2
Variable Definition and Summary: Contingent Valuation (Benefit)

Category and Description	Name	Unit	Observations	Mean (Std. Dev.)	Min.–Max.
Dichotomous choice					
Was first bid accepted?	Accepted 1	Binary	488	0.51 (0.5)	0–1
Was second bid accepted?	Accepted 2	Binary	488	0.44 (0.5)	0–1
Price offered in the first bid	Bid 1	R$	488	186.02 (158.32)	20–500
Price offered in the second bid	Bid 2	R$	488	172.58 (138.44)	10–600
Respiratory illness					
Duration of respiratory illness episode	Duration	Days	488	8.97 (8.41)	1–90
Number of symptoms caused by the respiratory illness episode	Symptoms	Count	488	10.32 (4.07)	1–20
Did the illness cause extreme pain/ discomfort? (=1 if yes)	Pain	Binary	488	0.18 (0.38)	0–1
Patient					
Age	Age	Years	488	29.95 (23.97)	0.33–89
Age squared	Age sq	Years^2	488	1,470.79 (1,835.33)	0.11–7,921
Gender (=1 for females)	Female	Binary	488	0.49 (0.5)	0–1
Smoke?	Smoke	Binary	488	0.13 (0.34)	0–1
Had ever had a serious lung disease?	Lung	Binary	488	0.36 (0.48)	0–1
Budget owner					
Years of schooling	Educ	Years	488	7.82 (4.32)	0–15
Household					
Logarithm of total household income	Income	R$/month	488	7.51 (0.76)	4.61–10.17
Total household income	Income	R$/month	488	2,416.89 (2,184.31)	100–26,000
Number of dwellers	Dwellers	Count	488	3.95 (1.84)	1–15

a low rate according to the literature on contingent valuation in health (see, e.g., Brandt, Lavín, and Hanemann 2012; Gyrd-Hansen 2016; Ready, Navrud, and Dubourg 2001).

Theoretical Validity: WTP Function

All econometric models estimated positive and significant coefficients for the number of symptoms and income, attesting to theoretical validity (Table 3). As expected, WTP to avoid illnesses for children (age < 18) was higher, due to their higher susceptibility (Alberini et al. 2010; Do Carmo, Alves, and Hacon 2013); however, the elderly (age > 65), whose respiratory health is also more vulnerable (Nunes, Ignotti, and Hacon 2013; Do Carmo, Alves, and Hacon 2013), did not have higher WTP.

Mean WTP in the single-bounded econometric models was R$ 192 and R$ 215 (8% and 9% of average income), with a range of R$ 161 to R$ 248. Mean WTP was smaller in the double-bounded models (omitted), providing further evidence for previously observed downward bias (Carson and Groves 2007). We have based our cost-benefit analysis on the lower mean from the single-bounded model (R$ 192, equivalent to $59). This value corresponds to 85% and 153% of the averages for, respectively, the upper and lower bound of estimates for the WTP to avoid respiratory illness informed by six previous studies, four of them referring to influenza vaccination (Prosser et al. 2005; Araña and Léon 2002; Hou et al. 2014; Low et al. 2017; Ortiz et al. 2011; Ara and Tekeşin 2016).

Cost of Fire Reduction

Description of Survey Data

The sample was representative of the SR, as shown in Appendix A, Section 5.2. Smallholdings were 55 ha on average. Cattle ranching was popular, as 85% of respondents had pasture and 81% had cattle as one of their main farming goals. Also, most respondents (86%) were integrated with markets: mainly cattle and dairy markets. Only four protest responses were detected. In all of these protest responses, respondents did not believe the expanded support for mechanization would be delivered by the government. Eleven additional respondents were excluded due to inter-

Table 3
Estimated WTP Function

	Linear	Boot Linear	Logit	Probit
Duration	0.2864 (1.7405)	0.2864 (2.1080)	0.0007 (0.0134)	0.0013 (0.0079)
Symptoms	12.8858*** (3.9285)	12.8858** (5.0637)	0.1020+ (0.0307)	0.0590*** (0.0180)
Pain	–2.1158 (39.7722)	–2.1158 (40.1325)	–0.010 (0.3043)	–0.009 (0.1796)
Age	–10.9627+ (2.4807)	–10.9627+ (2.7376)	–0.085+ (0.0190)	–0.050+ (0.0110)
Age sq	0.1145+ (0.0320)	0.1145*** (0.0357)	0.0008+ (0.0002)	0.0005+ (0.0001)
Female	–59.2078** (29.1913)	–59.2078** (28.4854)	–0.413* (0.2231)	–0.271** (0.1314)
Smoke	62.9024 (44.8000)	62.9024 (43.9651)	0.4995 (0.3475)	0.2883 (0.2043)
Lung	–31.6486 (30.7362)	–31.6486 (32.9669)	–0.263 (0.2384)	–0.145 (0.1405)
Educ	–0.1943 (4.0287)	–0.1943 (4.2517)	–0.002 (0.0316)	–0.000 (0.0185)
Income	94.1601+ (22.3625)	94.1601+ (25.6528)	0.7203+ (0.1742)	0.4316+ (0.1002)
Dwellers	–17.3702** (8.1898)	–17.3702 (10.6422)	–0.137** (0.0626)	–0.079** (0.0370)
Intercept	–388.0523** (159.0274)	–388.0523** (174.3191)	–2.938** (1.2239)	–1.778** (0.7131)
Bid			–0.007+ (0.0008)	–0.004+ (0.0004)
Intercept	218.1371+ (21.6910)	218.1371+ (24.1848)		
N	488	488	488	488
Chi^2	43.1204	38.3706		
p	0	0.0001		
Log-likelihood	–256.4987	–256.4987	–256.2068	–256.4987
Pseudo R^2	9.65%	9.65%	24.2%	0.2028
WTP:mean [CI]	192 [165:220]	192 [161:223]	215 [191:247]	215 [192:248]

Note: Standard errors in parentheses. Willingness to pay (WTP) could not be bounded by double-bounded probit, log-logistic, log-normal, and Weibull disturbance distributions. Boot, bootstrapped; CI, confidence interval.
* $p < 10\%$; ** $p < 5\%$; *** $p < 1\%$; + $p < 0.1\%$.

viewer mistakes. In the end, 207 respondents were considered for econometric modeling.

Over half of respondents (59%) thought that increased support for mechanization was unlikely. However, respondents showed greater belief in the command-and-control component of the amendments: 68% thought they were highly likely to be fined for exceeding burning limits. As a consequence, the policy scenarios presented in the survey were moderately or highly credible for 39% of respondents. This reflected the fact that most respondents did not have access to subsidized tractors (bulldozers and harrowing tractors had been received by only, respectively, 10% and 18%). Results also partly reflected the national sociopolitical environment prevailing during interviews (in May 2017), as the coupled political and economic crises—together with corruption scandals—made government appear untrustworthy (Kahn, Vásquez, and Rezende 2017).

Econometric Analysis of Survey Data

Variables' definitions and summaries are found in Table 4. In the four models estimated (Table 5), all attributes are significant and present the expected signs. There is evidence that respondents preferred policy options that gave them the right to burn a larger area. Heterogeneity of preferences is significant for all attributes except for harrowing tractor.

In all models the coefficient capturing preference for the status quo is highly significant and negative, revealing that the average respondent was an enthusiast of policy change. The estimate for the mean WTP for deviating from the status quo is R$ 39,081 ($11,988) in the MixWTP model, dominating all other attributes in the WTA formula. For instance, the mean WTP for the largest level of the most valued attribute, six additional bulldozer tractor hours per year, is R$ 7,944 ($2,437). This large effect[14] seems to stem from re-

[14] It should be stated that a large absolute marginal value for the SQ, compared with the other attributes, is not an uncommon finding. In de Valck et al.'s (2017) discrete choice experiment about nature restoration, the alternative specific constant capturing the non-SQ alternative had a larger estimated coefficient than most other attributes in some of the models estimated. In particular, in three models (table 4, Lovenhoek class 1; tables 5 and 6, Drogengoed, class 2), the ratio of non-SQ dummy coefficients and the average across coefficients of all attributes (except for price) was between

Table 4
Variable Definitions and Summary: Discrete Choice Experiment (Cost)

Category and Description	Name	Unit	Observations	Mean (Std. Dev.)	Min.–Max.
Dependent variable					
Chosen alternative	Choice	Binary	3,726	0.33 (0.47)	0–1
Attribute					
Allowed annual burning limit	Limit	Hectare/year	3,726	0.5 (0.41)	0–1
Increase in bulldozer tractor hours	Bulldozer	Tractor-hour/year	3,726	2.02 (2.46)	0–6
Increase in harrowing tractor hours	Harrowing	Tractor-hour/year	3,726	2.37 (3.27)	0–8
Reduction in tractor arrival delay	Delay	Binary	3,726	0.67 (0.47)	0–1
Compensation payment	Payment	R$/year	3,726	1,911.3 (2273.69)	0–6,000
Respondent[a]					
Used fire?	D_fire	Binary	3,726	0.45 (0.5)	0–1
Years of schooling	Educ	Year	3,726	5.36 (4.35)	0–15
Smallholding area	Area	Hectare	3,726	54.56 (47.04)	0.2–400
Area 1[b]	Area_1	Binary	3,726	0.39 (0.49)	0–1
Area 2	Area_2	Binary	3,726	0.32 (0.47)	0–1
Area 3	Area_3	Binary	3,726	0.11 (0.31)	0–1
Option					
Status quo dummy	SQ	Binary	3,726	0.33 (0.47)	0–1

Note: With 207 respondents, six choice sets per respondents, and three alternatives per choice set, number of observations totaled 3,726.
[a] Respondent variables are multiplied by the status quo dummy.
[b] Interviews took place in four smallholder clusters (Appendix A, Section 5).

spondents' interpretation that rejecting the status quo meant opting for increased access to insufficiently supplied public services that are complementary to tractors. For instance, many respondents complained about the lack of paved roads. Additionally, most respondents were poor and lived in remote lands and were therefore deprived of essential public services. These two reasons together explain the strong preference for non–status quo alternatives, which is also attested to by the low share of the status quo in the observed choices (10%). Another important result is that delayed tractor arrival caused a welfare loss of nearly the same magnitude, on average, as reducing the burning limit to zero.

Corrections in WTA Calculations

The empirical or estimated version of the WTA formula should include only the variables that significantly influenced respondents' choice as, according to theory, those should be the actual components of indirect utility functions (Bekker-Grobb, Ryan, and Gerard 2012). Few respondent-specific variables are significant, and this occurs only in two of the models estimated, the CLogit and GMNL. However, the CLogit model ignores the preference heterogeneity detected by other models, and the GMNL could not be estimated in WTP-space, as the only routine available (Hole 2011) continuously failed to start estimation—with no clear reason and without reporting an informative error message. Case-specific covariates were thus disregarded. For conservativeness, the status quo coefficient was not included in WTA estimates, as it strongly contributed to reduced compliance costs. A second reason for this exclusion was that the status quo coefficient seemed to capture public services that cannot be changed by the fire-for-tractors policy alone, but require complementary policies. Notwithstanding, other factors unrelated to attributes of the policy of interest (such as road paving) had to remain constant for an accurate measure of compensating variation to be obtained.

The final formula for estimating WTA the pth policy is

362% and 412%. This ratio was 185% in Christensen et al.'s (2011) discrete choice experiment on preferences for agrienvironmental contracts. The same ratio was 232% in Van den Broek et al.'s (2017) discrete choice experiment on selling contract choice involving African smallholders, and 273% for the Polish residents that Ščasný et al. (2017) confronted with choices on climate mitigation policy. In our study, the ratio attained the levels of 352% and 377%, respectively, in the preference-space and WTP-space mixed logit models, which is within the range of previous studies.

Table 5
Estimation Results

	CLogit	Mixed	GMNL	MixWTP[a]
Limit	0.44617+ (0.13276)	1.24154+ (0.36669)	1.33062*** (0.49189)	5,432.01822*** (1,981.14727)
Bulldozer	0.14858+ (0.0159)	0.26692+ (0.04643)	0.33431+ (0.06031)	1,324.49786+ (235.82751)
Harrowing	0.08879+ (0.0109)	0.1526+ (0.02812)	0.20401+ (0.03926)	734.55762+ (130.20218)
Delay	–0.65892+ (0.06742)	–1.08622+ (0.17)	–1.42691+ (0.25929)	–5,248.99399+ (1,027.39094)
Payment	0.00012+ (0.00002)	0.00022+ (0.00006)	0.00028+ (0.00006)	–8.58219+ (0.16791)
D_fire	0.57167** (0.28693)	0.64626 (1.73252)	1.45576*** (0.50788)	–574.63444 (8,931.14243)
Educ	0.03035 (0.0325)	–0.05822 (0.18299)	–0.24933+ (0.06878)	22.20909 (507.20357)
Area	0.00235* (0.00123)	0.01962 (0.01917)	–0.00623 (0.0056)	19.65906 (50.4571)
Area_1	0.19425 (0.45686)	0.93495 (1.73787)	–0.82535 (0.7398)	3,581.57706 (7,722.39327)
Area_2	0.77589* (0.4452)	1.1212 (1.47461)	–3.01359** (1.20492)	466.11207 (7,940.09259)
Area_3	0.09792 (0.6094)	–0.7897 (2.56203)	–1.7801* (0.95081)	–4,486.1399 (13,246.71216)
SQ	–2.18288+ (0.55517)	–6.54497** (2.92675)	–6.01196+ (1.62339)	–39,081.24101*** (14,232.33379)
SD				
Limit		–3.94151*** (1.37)	–5.15356+ (1.08991)	17,629.74053+ (3,200.26731)
Bulldozer		0.02719 (0.16699)	–0.13336+ (0.03641)	222.08299 (233.10439)
Harrowing		0.04893 (0.0911)	0.08038 (0.05825)	–205.91062 (134.75699)
Delay		–1.22169+ (0.32445)	1.49485+ (0.36895)	–5,551.72353+ (1,151.8322)
Payment		0.00026*** (0.00009)	0.00029+ (0.00009)	0.27029** (0.125)
SQ		4.1419** (1.97804)	5.42485+ (1.29962)	–28,282.38224*** (5,871.24625)
N	3,726	3,726	3,726	3,726
Chi^2	556.64	166.54	59.48	9,754.79
p	0	0	0	0
Log-likelihood	–882.89	–753.75	–738.39	–748.59
Pseudo R^2	0.15	0.22	0.24	0.21

Note: Standard errors (in parentheses) are robust for heteroskedasticity and autocorrelation.
[a] In the MixWTP model, estimates for "payment" correspond to the natural logarithm of payment (Hole 2011).
* $p < 10\%$; ** $p < 5\%$; *** $p < 1\%$; + $p < 0.1\%$.

$$\widehat{WTA}(p,i) = \hat{\mu}_{0i}(L_0 - L_p) - \hat{\mu}_{1i}bull'_p - \hat{\mu}_{2i}harr'_p - \hat{\mu}_{3i}del'_p, \quad [6]$$

where L_0 = 1 ha/farm-year is the status quo burning limit, and coefficients' estimates come from a MixWTP model.

Estimates of Population Sizes

Marginal effects were significant for the two stages (details in Appendix A, Section 3.2 and results in Appendix B, Tables B2 and B3). In the first stage, the significance was clearly due to the coincidence of peaks in the AOD and fire time series (see Appendix A, Figure A2), a pattern also found in previous studies (Mishra et al. 2015; Koren, Remer, and Longo 2007; Bevan et al. 2009). The time series were stationary, so the correlation captured by coefficient estimates was not spurious. In the second stage, AOD had a significant effect (1) on the total count of respiratory illnesses, (2) on all age group counts, and also (3) on all illness groups except for pneumonia.

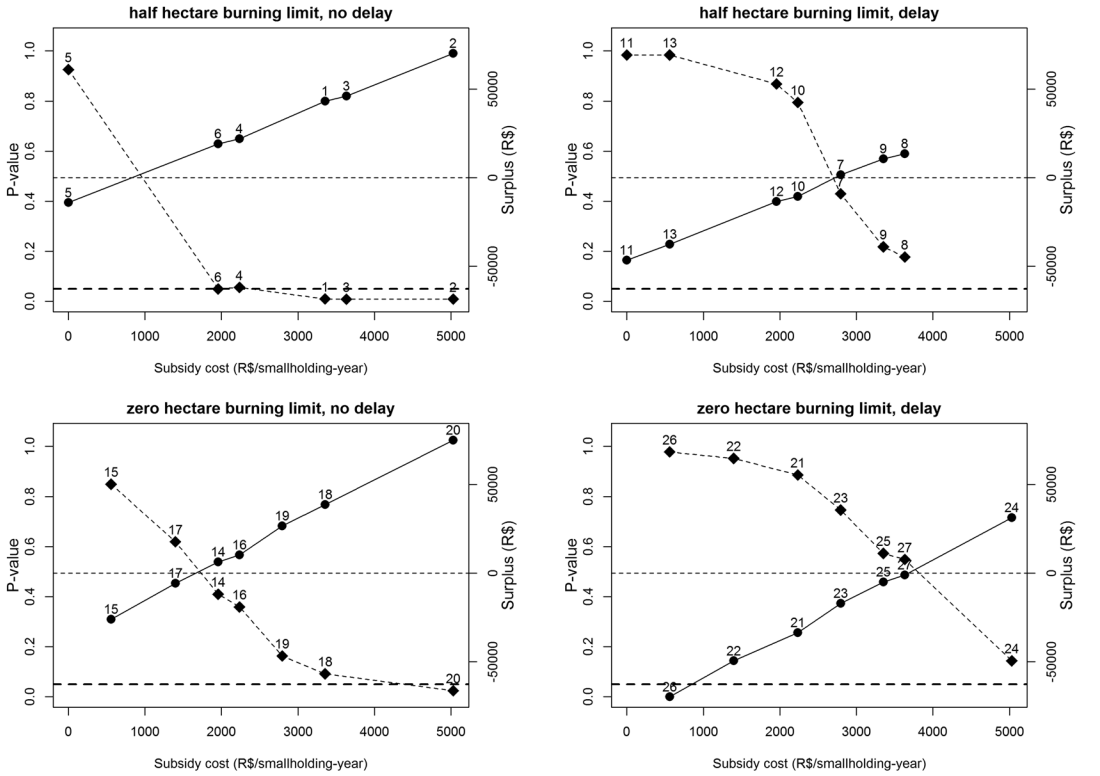
On average, reducing fire use at one smallholding benefited less than one urban dweller. The elimination of one fire detection avoided 2.64 illness cases, but reducing the burned area of the average smallholding to new limits (i.e., to half or zero hectare) eliminated a maximum of 0.08 point fire detections. Avoided illnesses equaled 8% of total illnesses observed in 2016. Under the scenario with a half-hectare burning limit, the worse-off population was 6,092 and the better-off was 1,247. For a zero burning limit, the estimates were 8,197 and 1,393, respectively.

Efficiency of Policies

In the SHK test, a decision about a policy's efficiency is based upon the test's *p*-value and the significance level. The latter is the probability of claiming an inefficient policy amendment is efficient; an error that is commonly tolerated in no more than 5% of cases. The *p*-value is based not only on the point estimate, but also on the empirical distribution

Figure 1

Efficiency Assessment of Policy Amendments as Function of Attribute Levels, Point Estimate for Surplus *(solid line)*, and *p*-Value of Statistical Hicks-Kaldor Test *(dotted line)*



of the underlying statistic. This explains why the SHK test captured different responses to attribute changes by different smallholders, as revealed by the nonlinear relationship between the test's *p*-value and attributes' levels. In contrast, a linear pattern prevailed for the point estimate.[15] This is visible in Figure 1,[16] which depicts the efficiency level of amendments grouped by burning limit and delay, and describes the central result of the paper.

First, the figure shows that extra subsidized tractor hours were not sufficient, alone, to make amendments efficient. Indeed, the effect of total tractor hours, expressed in terms of its cost for government, was counteracted by the effect of delayed arrival and a zero burning limit, characteristics that proved to be more influential on efficiency. This is clear from the patterns connecting the *p*-value and subsidy cost. Taking the half-hectare policy (amendment) without delay as a basis, both the half-hectare policy with delay and zero-hectare policy without delay show less abrupt decay patterns, which terminate with only one or no amendment below the significance level.

Second, it is also salient that, besides making amendments inefficient, delay altered the pattern of *p*-value decay, thus changing the benefit provided by extra tractor hours. In the two graphs with delay, an increase in tractor hours from the minimum level to R$ 2,000

[15] The willingness of the *i*th respondent to accept compensation for the *p*th policy was estimated as $\widehat{WTA}(p,i) = \hat{\mu}_{0i}(L_0 - L_p) - \hat{\mu}_{1i}bull'_p - \hat{\mu}_{2i}harr'_p - \hat{\mu}_{3i}del'_p = \beta_i x_p$. The surplus for the *p*th policy was $\hat{\Omega}_B\overline{B} - \hat{\Omega}_B \frac{1}{N_C}\sum_{i=1}^{N_C}\beta_i x_p = \hat{\Omega}_B\overline{B} - \hat{\Omega}_B\overline{\beta}x_p$, a linear function of x_p.

[16] In Figure 1, the upper horizontal line indicates zero surplus, and the lower horizontal line indicates the 5% significance level. The dots on the surplus and on the *p*-value lines are the policy amendments evaluated (numbers match those of Table 6). Tractor hours are expressed as direct cost for the government, hereafter "subsidy cost" (Appendix A, Section 2.3).

had a smaller effect on the p-value compared to graphs without delay (see Figure 1).

Third, sampling error impacted the assessment of efficiency. In all graphs there are more amendments with positive point estimates than with p-values below 5%. In fact, based on the point estimate, 14 amendments are classified as efficient. However, only 5 amendments are classified as efficient based on the SHK test (Figure 1, Table 6). Therefore, the critical value threshold effectively makes the SHK test more demanding as an efficiency judgment. This is clear from the fact that only the four amendments with highest estimated surplus exceed the statistical threshold by a nonnegligible magnitude (Table 6). Also, based on empirical distributions, nonefficient amendments have a statistically larger absolute value for the ratio of standard deviation/mean of surplus (p-value < 0.01%), showing that the threshold is proportional to sampling error.

It should be added that the nine "false" efficient amendments provided small extra subsidies in exchange for big fire reduction sacrifices and were bad "deals" for many smallholders—either tractors were delivered with delay or a zero burning limit was imposed. Avoiding the mistake of prescribing these detrimental amendments is an important practical implication of the SHK test.

Finally, the amendments that are efficient increased tractor hours to the maximum level for at least one of the tractors, and delivered them without delay to smallholders (Table 6). The only exception is amendment 6, which is borderline efficient. Only one efficient policy had a zero burning limit, but it offered the largest number of extra tractor hours without any delay.

Let "value for money" be understood as the cash compensation that is saved by each R$ 1 expended on "in-kind" subsidy. In this case, compensation targets the losses imposed by reduced burning limits, and an in-kind subsidy is provided in the form of extra tractor hours. The benchmark level of value for money is R$ 1, and it is not met by efficient amendments, except for number 6, which merely achieves the cutting edge of the SHK test and has the smallest surplus among the efficient amendments (Table 6); for amendment 6, the point estimate for surplus exceeds the critical value by 0.001 percentage point. Two implications follow. First, providing a moderate increase in tractor hours, with on-time arrival and a half hectare burning limit, was exactly enough to meet the efficiency threshold, which is the case of amendment 6. Under the latter policy change, a 61% saving was made compared to the amendment that maximized surplus. Second, there is a trade-off between social welfare maximization and public expenditure minimization, so policy makers prioritizing the latter should compensate smallholders for a lower burning limit with cash. In contrast, it is advisable that "welfarist" policy makers compensate with increased tractor hours—with the amount and punctuality at the efficient level.

The largest surplus for half and zero burning limits was broken down into aggregate WTP and WTA at the level of SR sectors. This revealed that, for both worse-off and better-off populations, the east of the SR should be a top priority for policy amendments (see Table B4 in Appendix B). However, it must be highlighted that, as sub-SR aggregates suggest, the harsher fire constraints evaluated in this study are not justified by the health benefit, but by the tractor benefit alone. For the five efficient amendments, the latter exceeds the former at least 140-fold.

To further assess the robustness of the efficiency results, the influence of the sampling error inherent in the estimate of the size of the better-off population was also evaluated. Two extreme values for the estimate were calculated, based on the minimum and maximum values for the coefficient measuring the effect of smoke on respiratory illnesses (namely, 0.002 and 0.123, see Section 3; only models with 100% significance rate were considered). The outcome is that the efficiency of the amendments did not qualitatively differ between the extreme estimates for the population size and the "baseline" case detailed in the previous paragraphs (which built on a 0.006 coefficient). In fact, the difference in the extreme estimates was an additional efficient (maximum estimate) or inefficient (minimum estimate) amendment. However, the amendments with different statuses present the same characteristics as three of the four policies that

Table 6

Detailed Efficiency Assessment of Policy Amendments

Policy Amendment	Tractor Hours (bull, harr)	Subsidy Cost (R$/farm-year)	Burning Limit	Delay (Y/N)	Surplus	HK Threshold	Efficient?	Welfare Loss: Fire Reduction	Welfare Gain: Tractor Hours	Value for Money	Surplus/HK Threshold (%)
1	6, 2	3,353	0.5	No	43,142	27,252	Yes	2,344	–9,386	0.70	158.3
2	6, 8	5,030	0.5	No	70,118	38,746	Yes	2,344	–13,814	0.47	181.0
3	3, 8	3,633	0.5	No	46,025	29,340	Yes	2,344	–9,859	0.65	156.9
4	0, 8	2,236	0.5	No	21,931	22,037	No	2,344	–5,904	1.05	99.5
5	0, 0	—	0.5	No	–14,037	14,423	No	2,344	—	NA	–97.3
6	3, 2	1,956	0.5	No	19,048	19,039	Yes	2,344	–5,431	1.20	100.0
7	6, 0	2,794	0.5	Yes	1,656	24,783	No	2,344	–7,910	0.84	6.7
8	3, 8	3,633	0.5	Yes	13,531	23,172	No	2,344	–9,859	0.65	58.4
9	6, 2	3,353	0.5	Yes	10,648	24,905	No	2,344	–9,386	0.70	42.8
10	0, 8	2,236	0.5	Yes	–10,563	18,444	No	2,344	–5,904	1.05	–57.3
11	0, 0	—	0.5	Yes	–46,531	20,909	No	2,344	—	NA	–222.5
12	3, 2	1,956	0.5	Yes	–13,445	20,494	No	2,344	–5,431	1.20	–65.6
13	0, 2	559	0.5	Yes	–37,539	19,961	No	2,344	–1,476	2.64	–188.1
14	3, 2	1,956	0	No	6,365	39,492	No	4,687	–5,431	2.40	16.1
15	0, 2	559	0	No	–26,054	38,606	No	4,687	–1,476	2.64	–67.5
16	0, 8	2,236	0	No	10,243	37,829	No	4,687	–5,904	2.10	27.1
17	3, 0	1,397	0	No	–5,734	37,848	No	4,687	–3,955	2.83	–15.2
18	6, 2	3,353	0	No	38,784	49,730	No	4,687	–9,386	1.40	78.0
19	6, 0	2,794	0	No	26,685	45,015	No	4,687	–7,910	1.68	59.3
20	6, 8	5,030	0	No	75,081	57,120	Yes	4,687	–13,814	0.93	131.4
21	0, 8	2,236	0	Yes	–33,478	37,464	No	4,687	–5,904	2.10	–89.4
22	3, 0	1,397	0	Yes	–49,455	42,081	No	4,687	–3,955	2.83	–117.5
23	6, 0	2,794	0	Yes	–17,036	47,362	No	4,687	–7,910	1.68	–36.0
24	6, 8	5,030	0	Yes	31,360	51,976	No	4,687	–13,814	0.93	60.3
25	6, 2	3,353	0	Yes	–4,937	46,911	No	4,687	–9,386	1.40	–10.5
26	0, 2	559	0	Yes	–69,775	41,930	No	4,687	–1,476	2.64	–166.4
27	3, 8	3,633	0	Yes	–1,059	42,200	No	4,687	–9,859	1.29	–2.5

Note: HK, Hicks-Kaldor; NA, not available.

prove efficient in the baseline: a half hectare reduction in the burning limit, no delay in tractor arrival, and a moderate increase in tractor hours. In addition, the amendments whose efficiency changed in the robustness check were "borderline" near-efficient cases in the baseline (amendments 4 and 6 of Table 6).

5. Discussion and Conclusion

The present paper seeks to offer both policy-relevant and methodologically sound contributions. With respect to the former, the economic performance of a policy to reduce the health impacts generated by agricultural fires in the Amazon was evaluated. The policy combines cap-and-surveillance of fire use and in-kind subsidies in the form of tractors—the latter being a substitute good for fire highly desired by small-scale fire users. Amending this policy to intensify its "stick" and "carrot" components proved beneficial for the respiratory health of urban dwellers and the welfare of rural fire-dependent smallholders, provided that reductions in the right to burn were compensated for with an increase in mechanization support.

Turning to consider the methodological contribution of the paper, the proposed statistical version of the Hicks-Kaldor test proved an effective antidote to address the sampling error introduced by preference heterogeneity. This test avoided an overstatement of the number of efficient policy amendments that would triple the correct number.

Our analysis also considered spatial heterogeneity. The findings identified the east of the SR as the top priority location for policy intervention. The main factor that drove this result was the number of smallholders burning above the amended area limits, 62% of them located in the east. This explains not only the east's larger aggregate WTA but also larger aggregate WTP values; the latter was estimated as a product of the factor already mentioned, which dominated the estimate, and two factors with lower or null variation within the study region (the marginal impact of fires on health and average WTP, respectively). Most of the smallholdings at the east of the study region belong to Acre's largest agrarian settlement, Pedro Peixoto, which occupies 57% of the eastern SR area. This result is in line with previous studies that detected high levels of deforestation and fire activity in the Peixoto settlement (Ávila and Wadt 2015; D'Oliveira and Braz 2006; IPEA 2014). Anderson et al. (2017) have also found evidence that fires detected in Acre state are mainly concentrated in agrarian settlements, and also that the eastern areas in Acre state (which contain the eastern part of the present SR) have the highest level and growth rate of fire frequency.

It was also evidenced that the efficient intensification of the fire-for-tractors policy is conditional on the fine-tuning of the policy's attributes. In particular, two main conditions must be met. First, total hours of the bulldozer and harrowing tractors need to be increased to a minimum of five hours per smallholding per year. Second, it is vital that tractors arrive on time at each smallholding, which requires punctuality and also that arrival be scheduled for the best period of the year for smallholders (July to September for the most of the sample). There are two ways to meet this second requirement. The first is to increase the mechanization services currently offered by the government. This increase would require a large investment, given the current level of delays[17] and the constraints imposed by low-quality roads.[18] The second is to schedule smallholders' individual farming agendas to avoid demand peaks above supply capacity. This scheduling would require consensus among smallholders and government technicians.

The analysis also demonstrates that if policy makers wish to minimize expenditure, they should offer cash compensation for a reduced legal burning limit. However, if the policy objective is to maximize smallholders' welfare, then efficient in-kind compensation, via subsidized mechanization, will be optimal. This second approach seems more con-

[17] Average delay was 40 days in the second pilot wave and in the full-fledged survey. The number of respondents that faced delay among those that received tractors was, in the second pilot wave, 10 out of 15 respondents with tractors (66%) and, in the full-fledged survey, 26 out of 35 respondents with tractors (74%).

[18] Information gathered during visit to Safra, December 2016.

sistent with current policy. For example, the withdrawal of the full ban on agricultural fires by Acre government, and its substitution with a fire-for-tractors approach, was defended on the grounds of protecting smallholders' subsistence agriculture (Acre 2013). Additionally, many existing policies specifically target improvements in smallholders' living standards through changes to land management.[19] In keeping with the apparent aims of current policy, amendments that qualified as efficient in this analysis did not reduce smallholders' welfare.

Mendoza-Escalante, Börner, and Hedden-Dunkhorst (2003) also detected a preference for tractor use among eastern Amazon smallholders. More precisely, those authors found a preference for replacing fires with mechanized mulching. This finding is consistent with the results of the present paper. Mburu et al. (2007) also found that replacing fire use with mulching tractors yielded a positive net benefit for smallholders, but only if cropping cycles and varieties were also altered. This result is similar to our finding that (re-)scheduling will be necessary to avoid tractor delay. In both cases, it will be necessary for smallholders to adapt their farming routine. Some barriers to reducing agricultural fire use will require government support to remove; these include the high costs of tractor and complementary inputs (Börner et al. 2007, Kato et al. 1999). To remove these barriers, subsidies—such as the additional tractor-hours evaluated here—may provide potential solutions.

It seems important to close by considering the very significant health benefits that the economically efficient amendments bestow upon the majority urban and rural population in terms of reduced pollution levels. During one-third of every year, the pollution level in the study region remains above the level of mega cities such as São Paulo and London, exposing over 400,000 inhabitants to health risks. Policies such as those considered in this paper can significantly contribute to addressing this problem.

Finally it is worth noting that the assessments presented in the present paper remain incomplete. Additional benefits of a move away from fire-based agriculture include reductions in greenhouse gas emissions, and decreases in both the degradation of forests and the loss of biodiversity (Anderson et al. 2015). Extension of our analysis to embrace these benefits is likely to justify the further acceleration of such policies and accordingly is a focus of our ongoing research.

[19] Personal communication with representatives of the Acre state agricultural support department (Seaprof), and conversations with the heads of the institutions participating in the farm certification and sustainable development programs, January 2017.

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