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Testing transferability of willingness to pay for forest fire prevention among three states of California, Florida and Montana

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Abstract

The equivalency of willingness to pay between the states of California, Florida and Montana is tested. Residents in California, Florida and Montana have an average willingness to pay of \$417, \$305, and \$382 for prescribed burning program, and \$403, \$230, and \$208 for mechanical fire fuel reduction program, respectively. Due to wide confidence intervals, household WTP in the three states are not statistically different. Over all tests, there is mixed evidence on transferability, but California and Montana WTP are similar to each other for prescribed burning and Florida and Montana have similar values for the mechanical fuel reduction.

JEL classification: Q260; Q230
Keywords: Benefits transfer; Contingent valuation; Mechanical fuel reduction; Prescribed burning

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Introduction

On August 20, 2002 President George W. Bush approved the Healthy Forests Initiative aiming for restoration of health forests and rangelands in the western United States. To restore the health of forests and rangelands, President Bush is seeking: first to improve procedures for developing and implementing fuel treatment and forest restoration projects in collaboration with local government, second to develop guidance for weighing the short-term risks against the long-term benefits of fuel treatment and restoration projects, third to develop guidance to ensure consistent National Environment Policy Act (NEPA) procedure for fuel treatment and restoration activities (Bush, 2002).

The two main fuel treatment methods considered in the Healthy Forest Initiative are: the prescribed burning and the mechanical fire fuels reduction. The prescribed burning method is defined as the controlled application of fire to existing naturally occurring fuels under specified environmental conditions following appropriate precautionary measure (Florida Division of Forestry, 2000). The mechanical fire fuel reduction method consists of mechanically removing smaller trees and vegetation. This mechanical fuel reduction method is especially effective at lowering the height of vegetation, which reduces the ability of fire to climb from the ground to the top or crown of the trees.

There are currently not available sources of valuation information or market signals that reveal the demand for these fire fuel reduction programs, especially public lands in many of the forested states in the US. Providing this type of information would allow the forest managers, and policy makers, to determine which states have the highest values for the prescribed burning and mechanical fire fuel reduction. This may help in allocating the scare resources for fire prevention programs.

As with many non-marketed natural resources, valuation of the protection of forest health and public forests is problematic. In part, this is due to the fact that protection of forest health includes both public recreation use values, downstream water quality, protection of forest dependent wildlife, and existence values from knowing that these forests are in good ecological condition for current and future generations (i.e., passive use values). Taken together these use and passive use values represent the total economic value of forest health (Randall and Stoll, 1983). Because total economic value contains both use and unobservable passive use values, a stated preference method such as contingent valuation (CVM) is needed to elicit total economic value (Randall and Stoll, 1983). The contingent valuation method involves developing a simulated or hypothetical market or referendum to elicit willingness to pay (WTP) or willingness to accept (Mitchell and Carson, 1989). While reliance on statements regarding WTP are viewed with some skepticism by some economists, the method has proven reliable in test–retest studies (Loomis, 1989). Contingent valuation has been used in past studies of the values of forest recreation and protecting public old growth forests for the spotted owl (Rubin et al., 1991), and for protecting old growth forests in Oregon from fire (Loomis and Gonzalez-Caban, 1997).

However, it would be very expensive to conduct surveys in all forested states in the US. This is a common problem, not only in forest valuation, but for water quality, recreation, etc. The limited ability to conduct site-specific studies for all public natural resources of policy interest has given rise to the field of benefit transfer (Boyle and Bergstrom, 1992). This field of study investigates the ability and accuracy of transferring benefit estimates from previous studies and applying them to new policy evaluations. The accuracy of benefit transfer is of great interest to resource managers as the possibility of accurate benefit transfers promises to reduce the time and expense of having to conduct original studies every time a new policy evaluation is performed.

The typical approach to testing the accuracy of benefit transfers is to compare original WTP estimates or WTP functions that have been estimated in one geographic location to what a benefit transfer would estimate as the value. For example, a typical test of the accuracy of benefit transfer might involve comparing an original study estimate of what households in Finland might pay for deer hunting to a benefit transfer for estimate of deer hunting based on a WTP function for deer hunting from Sweden.

Our study states of California (CA), Florida (FL) and Montana (MT) being located in the West Coast, East Coast and Northern Rocky Mountains, respectively, of the United States of America (US) provide a good opportunity to test for transferability of benefits of fire fuel reduction. Among these states there exist several differences including demographics (e.g., age and education), ecological differences in forest type, and of course the extent of wildfires. The residents in these three states may view wildfire reduction programs differently leading to the difference in how they value these programs. However, if there is some degree of similarity of WTP between these states, then forest managers may not have to develop state-specific estimates for much of the US.

The first objective of the study is to determine if differences exist in CVM survey response rates in California, Florida and Montana on two programs: prescribed burning and mechanical fire fuel reduction (hereafter RX and mechanical programs). The second objective is to compare the protest refusal to pay responses of people in California, Florida and Montana. Here we would like to find out the reasons why people place a zero value on the two programs. The third objective is to find out whether willingness to pay (hereafter WTP) of people in the three states for the two programs is affected by geographic differences or not. Finally, we test whether WTP per household is similar and the WTP functions are transferable between the three states or not. If the benefit estimates or WTP functions are not transferable, then surveys will have to be conducted in each geographic region to estimate the benefits of these two fire prevention methods.

Hypothesis tests on response rate and protest responses

Our survey modes involve an initial random digit phone call with a short (5 min) initial interview. Names and addresses of respondents are requested for mailing a survey book, followed by scheduling an in-depth interview (20 min) using the booklet. Thus, the first basis of comparison is whether people in three states of CA, FL and MT respond equally to the initial phone calls and follow through on the in-depth interview. The null hypothesis is that overall survey response rate (RR) to the CVM survey is independent of state residence:

$$\mathrm{H}_0:\ \mathrm{RR}^{\mathrm{CA}} = \mathrm{RR}^{\mathrm{FL}} = \mathrm{RR}^{\mathrm{MT}}. \tag{1}$$

A χ^2 test will be used for to test this hypothesis for the initial and in-depth interviews.

Responses to the WTP questions during the in-depth interview are the main focus of our analysis. Some refusals to pay are valid expression of zero WTP since they reflect lack of value for the good or low income (i.e., inability to pay). Other refusals to pay reflect protest against some features of the CVM scenario. The null hypothesis is that differences in protest and non-protest responses (PR) are the same among people in three states of California, Florida and Montana for the RX and mechanical programs:

$$\text{Null hypothesis for RX program}:\ \mathrm{H}_0: \mathrm{PR}^{\mathrm{CA}} = \mathrm{PR}^{\mathrm{FL}} = \mathrm{PR}^{\mathrm{MT}}, \tag{2}$$

$$\text{Null hypothesis for mechanical program}:\ \mathrm{H}_0: \mathrm{PR}^{\mathrm{CA}} = \mathrm{PR}^{\mathrm{FL}} = \mathrm{PR}^{\mathrm{MT}}. \tag{3}$$

The test of significance will be performed using a χ^2 test.

WTP model and related hypothesis tests

As recommended by the NOAA panel (Arrow et al. 1993), we used a voter referendum format to ask the willingness to pay question. According to Hanemann (1984), respondents evaluate the utility difference associated with the current program level versus paying some amount of money ($\$X$) for an increase in the program level. If the utility difference is positive for the program, the individual is believed to respond "yes". In terms of our specific empirical analysis we formulate the utility function of the respondent in terms of the relative level of the public good, the percentage reduction in acres of wildfire in the respondent's state. A utility function in which the consumer evaluates their utility based on the relative levels rather than absolute levels of the public good is often called a reference-dependent utility function (Hanemann, 1999). Use of percentage changes as the attribute of valuation rather than absolute acres is not uncommon in CVM and choice experiments. For example, in evaluating different alternative levels of nature conservation in Finland, Li et al. (2004) used percentage changes from the current amount of land preservation and percentage of the land area in Finland as the primary attribute levels. Thus, we formulate a simple utility function in which the individual receives utility from income (I) and the percentage reduction in forest fires in their state (F):

$$U = f(I, F). \tag{4}$$

Because the utility function is not completely observable by the researcher:

$$U = f(I_0, F_0) = V(I_0, F_0) + e_0, \tag{5}$$

where e is the unobservable part of utility that is considered by the research to be a random variable (Hanemann, 1984). In a dichotomous choice CVM willingness to pay question, the individual is offered a higher level of the public good in exchange for a reduction in income by the bid amount (\$$X$). In our case study, the increase in the public good (F) is forest fuel reduction program that will result in an increase in forest health from a 25% reduction in acres of forests burned in catastrophic wildfires (F_1). With this program the utility is

$$V(I_0 - \$X, F_1) + e_1. \tag{6}$$

A utility maximizing individual is believed to make their decision as to whether to respond "Yes, they would pay \$$X$" or "No they would not pay" by comparing the baseline utility (Eq. (5)) of forest health to the utility derived from the program (Eq. (6)). If the difference in utility (Eq. (6) minus Eq. (5)) is positive, the individual is predicted to state Yes they would pay.

If the utility difference is distributed logistically, a logit model can be used to estimate the parameters and allow for calculation of WTP (Hanemann, 1984).

To test for the effect of different states on willingness to pay responses, two tests were conducted. We can test whether the state of residence simply shifts the logit index function up or down by a state dummy variable or rotates the logit index by using an interaction term on the bid amount. In Eqs. (7) and (8) below, for each program we are subsuming Montana as the base case.

First we define the odds of voting for each of the fuel reduction programs is $A = P_i/(1 - P_i)$. Then the logistic regression equation for prescribed burning is

$$\begin{aligned}\mathrm{Ln}\,(\mathrm{A}) = \beta_0 + \beta_1\mathrm{Bid} + \beta_2\mathrm{FL} + \beta_3\mathrm{FL} * \mathrm{Bid} + \beta_4\mathrm{CA} + \beta_5\mathrm{CA} * \mathrm{Bid} \\ + \beta_6 X_6 + \ldots + \beta_n X_n + u_i.\end{aligned} \tag{7}$$

Similarly for the mechanical fire fuel reduction program:

$$\begin{aligned}\mathrm{Ln}\,(\mathrm{A}) = \beta_0 + \beta_1\mathrm{Bid} + \beta_2\mathrm{FL} + \beta_3\mathrm{FL} * \mathrm{Bid} + \beta_4\mathrm{CA} + \beta_5\mathrm{CA} * \mathrm{Bid} \\ + \beta_6 X_6 + \ldots + \beta_n X_n + u_i,\end{aligned} \tag{8}$$

where Bid- the dollar amount of bid the respondent is asked to pay; FL—Florida, CA—California are shift variables and they equal 1 for residents from the states of Florida and California, respectively, and 0 otherwise, FL∗Bid, CA∗Bid are interaction terms and u_i is the stochastic disturbance term with normal distribution with zero mean (Gujarati, 1997).

If Florida and California WTP functions are similar to Montana, then

$$\mathrm{H}_0: \beta_2 = 0;\ \beta_3 = 0;\ \beta_4 = 0;\ \beta_5 = 0. \tag{9}$$

The hypotheses are tested individually using t-statistics on β_2, β_3, β_4, β_5. However, it may be that we reject the null hypothesis of equality of California and Florida with Montana, but that California and Montana are similar. That is, $\beta_2 = \beta_3$ and $\beta_4 = \beta_5$ and none of these may be equal to zero. Thus, if we reject the null hypothesis of equality of California and Florida intercepts with Montana, and equivalently for two states bid slope interaction coefficients, then we will test for the equality of the California and Florida intercepts with each other, and the same for their bid slope interaction coefficients. This will be tested using a *t*-test for the equality of the coefficients.

A more general test is to evaluate whether one or more of logit coefficients in Eqs. (7) and (8) vary with states. To test this, we estimate logit models for each state.

For prescribed burning program:

$$\text{California,}\quad \mathrm{Ln(A)} = \alpha_0 + \alpha_1 \mathrm{Bid} + \alpha_2 X_2 + \alpha_3 X_3 + \cdots + \alpha_n X_n + u_j, \tag{10}$$

$$\text{Florida,}\quad \mathrm{Ln\,(A)} = \gamma_0 + \gamma_1 \mathrm{Bid} + \gamma_2 X_2 + \gamma_3 X_3 + \cdots + \gamma_n X_n + u_j, \tag{11}$$

$$\text{Montana,}\quad \mathrm{Ln\,(A)} = \delta_0 + \delta_1 \mathrm{Bid} + \delta_2 X_2 + \delta_3 X_3 + \cdots + \delta_n X_n + u_j, \tag{12}$$

The null hypotheses:

$$\mathrm{H}_0 : \alpha_0 = \gamma_0 = \delta_0;\ \alpha_1 = \gamma_1 = \delta_1;\ \alpha_2 = \gamma_2 = \delta_2;\ \cdots \alpha_n = \gamma_n = \delta_n. \tag{13}$$

Likelihood ratio test on these separate equations for the program will be conducted.

We do the same for mechanical fire fuel reduction program.

The willingness to pay is calculated using the formula proposed by Hanemann in 1989:

$$\text{Mean WTP}_{\mathrm{CA}} = \ln(1 + \exp(\alpha_0 + \alpha_2 X_2 + \alpha_3 X_3 + \ldots + \alpha_n X_n))/\mathrm{ABS}(\alpha_1) \tag{14}$$

for people in California and using respective logit model coefficients (ABS—absolute value). We use the same formula for $\mathrm{WTP}_{\mathrm{FL}}$, $\mathrm{WTP}_{\mathrm{MT}}$ with the state respective coefficients.

To test the state effects on willingness to pay, we compare mean WTP households across three states for each program with the null hypothesis:

$$\mathrm{H}_0 : \mathrm{WTP}_{\mathrm{CA}} = \mathrm{WTP}_{\mathrm{FL}} = \mathrm{WTP}_{\mathrm{MT}}, \tag{15}$$

with the state respective coefficients.

The null hypotheses state that WTP of California's residents are the same as that of Florida's residents and Montana's residents. The null hypotheses would be tested by whether the confidence intervals overlap or not.

Survey design

The survey booklet began by discussing large wildfires in three states in the year before the survey. It contained information and drawings contrasting wildfires and prescribed burning as part of the description of the expanded public forest fuel reduction program.

The following WTP elicitation question was used for prescribed burning program:

If the Expanded Prescribed Burning Program was undertaken in your county and state, it is expected to reduce the number of acres of wildfires from the current average of approximately AAAA acres each year to about AAA acres, for a 25% reduction.

Your Chance to Vote: Your share of the Expanded Prescribed Burning Program would cost your household $X... a year. If the Expanded Prescribed Burning Program were on the next ballot would you vote: In Favor ... Against...

A similar question also was used for mechanical fire fuel reduction program. As noted above, our null hypothesis of benefit transferability across states assumes the relevant measure of the benefits of the fuel reduction programs is the equivalent 25% reduction in wildfires across states, rather than the absolute acreage reduction (which varied across the three states). If respondents are focusing on the absolute acres of wildfire reduction, then we would not necessarily expect equal mean WTP per household in each of the three states, although the WTP functions might still be similar.

Ten bid amounts denoted $*X*, randomly varied across states for both program. These amounts for the mechanical fire fuel reduction are on average $10 higher than those of prescribed burning program. The bid amounts for prescribed burning were $10, $20, $30, $40, $60, $90, $120, $150, $250, and $350.

After the question on willingness to pay is asked, if a respondent indicated she or he would vote against the program, then they were asked an open-ended question: "Why did you vote this way?". The reasons obtained are content analyzed to classify answers by similar reasons given by the respondent. This open-ended response approach avoids having respondents fit themselves into pre-set protest categories or interviewers placing them into those categories. The final page of the booklet is the demographics section.

Data collection and survey mode

To obtain a representative sample in three states of California, Florida and Montana, a random digit dialing of the population was used. The use of random dialing assures that nearly all households are eligible to be interviewed, whether they have listed phone number or not. The surveys were conducted using a phone–mail–phone approach. The initial phone interview lasted about 5 min with questions focusing on the introduction of the survey purpose, assessing preliminary knowledge of respondents on fires and obtaining address to send the in-depth survey booklet. The individuals were asked to read the booklet prior scheduled date of phone interview.

Response rate analysis

There are two types of response rates to be examined: the first is screener response rate (or the first wave RR) and the second is in-depth interview response rate (or the second wave RR). The first wave RR is the percent of respondents from the total initial sample that has been contacted and those completed the initial interview. The percent of net sample completed in the in-depth interview is the second wave

Table 1. Response rates in California, Florida and Montana

	California		Florida		Montana	
	Persons	%	*Persons*	%	*Persons*	%
First wave						
Total initial sample	794		626		602	
Completed initials	328	41.3	534	85.3	407	67.6
χ^2 of first wave			69.89***			
Second wave						
Net sample of second interview	257		454		373	
Completed interviews	187	72.8	328	72.2	272	72.9
χ^2 of the second wave			0.008			
χ^2 critical at 5% and 1%			5.99 and 9.21			
Degree of freedom			2			

***Statistically significant at the 1% level.

response rate. The response rates in the first wave of interview in CA, FL and MT are 41.3%, 85.3% and 67.6%, respectively (Table 1). The χ^2 statistic of the first wave response rate ($\chi^2 = 69.89$) is significant at the level of 1%. Therefore, we can infer that there is a statistically significant difference among response rates across people in CA, FL and MT to the initial phone call. In the in-depth wave interview, the χ^2 statistic is not significant at the level of 1% and 5%. The response percentages in three states CA, FL and MT in the second wave are similar (72.8%, 72.2% and 72.9%, respectively). This means that response rates among people in CA, FL and MT are not significantly different for the CVM survey. Completing the more in-depth CVM interview, we could see increase in response rates of people in CA and MT (31.5% and 5.3%, respectively) compared to those rates of the first wave interview.

Refusal to pay analysis

The recording of open-ended statements after respondents voted "no" to a specific fuel treatment program allowed for identification of protest and non-protest votes. The reasons like opposition to all government programs, stating the program would not work, opposed to taxes, etc. are considered to be protest votes. Alternatively, reasons for the no votes by respondents such as the program is not worth the money or they cannot afford paying for the programs are the non-protest votes and show that the respondents are taking the contingent market seriously. Some of these non-protest refusals to pay may relate to the loss in utility from the smoke from prescribed burning outweighing the utility of wildfire reduction. As such, an alternative modeling approach would be to use a spike model of Kristrom (1997) as is demonstrated by Nahuelhual et al. 2004. The refusals to pay in our study are examined for each program: prescribed burning and mechanical fire fuel reduction.

Table 2 presents the responses categorized and identified as protest and non-protest. The calculated χ^2 of 3.368 (Table 2) with degree of freedom of 2 indicates no statistically significant difference among residents in three states CA, FL and MT in the pattern of protest and non-protest reasons for refusing to pay for prescribed burning program. We could say that the ratios between protest and non-protest refusal to pay responses of people among three states CA, FL and MT are equivalent to each other and independent of states. Further, the protest rates are quite low for the RX program.

Refusals to pay for the mechanical program were also categorized into protest and non-protest responses for three states CA, FL and MT. The χ^2 of 2.009 (Table 2) has been calculated for protest and non-protest responses of refusal to pay the bids. From calculated χ^2, it is obvious that there is no statistically significant difference among people in three states in the pattern of saying no to the bids for mechanical program. However, looking at Table 2 we could see more protests by respondents to the mechanical program than the prescribed fires. From 273 interviewees in MT, there were 36 households that were protests, and this number in FL was 21 households. Comparing this indicator of mechanical program to that of RX program, it could be deduced that people in three states CA, FL and MT are more supportive of the RX program as compared to mechanical program.

Table 2. Comparison of refusals to pay

	California		Florida		Montana	
	Persons	%	*Persons*	%	*Persons*	%
Prescribe burning						
Protest responses	8	4.3	14	4.3	17	6.2
Non protest responses	6	3.2	18	5.5	8	2.9
χ^2 calculated			3.368			
χ^2 critical at 5% and 1%			5.99 and 9.21			
Degree of freedom			2			
Mechanical						
Protest responses	13	6.9	21	6.4	36	13.2
Non protest responses	12	6.3	38	11.6	49	17.9
χ^2 calculated			2.009			
χ^2 critical at 5% and 1%			5.99 and 9.21			
Degree of freedom			2			

Statistical analysis of willingness to pay responses

Development of logistic regression began with building the initial model based on the selected variables (Table 3). To reflect the impact of geographic difference on probability of voting for the proposed fuels reduction programs, we include in the model two dummy variables: California and Florida. The state of Montana is subsumed as the base case. Besides these, we also include interaction terms of state variables and bid amount. The purpose of this is to test for equality of the impact of the bid amount in each state on probability of voting in favor of the program.

From Table 4, for prescribed burning, California and Florida state intercept variables are significantly different than zero (and hence from Montana) for the RX logit models with protest included and protests excluded. Thus, in term of our hypothesis regarding states, the state logit intercepts do shift up the logit functions by the values of coefficients for the RX program relative to the respondents in Montana. This suggests some differences between Montana and the two other states. However, the California and Florida intercept shifters are not statistically different from one another using a *t*-test of coefficient equality ($t = .236$ for the protest-included case) suggesting similarity between California and Florida in this regard. None of state bid interaction terms are significant at the 5% or 1% level. Thus the sensitivity to the bid amount is not statistically different than Montana residents. Thus overall, the null hypothesis of H_0: $\beta_2 = 0$ and $\beta_4 = 0$ is rejected for RX program, and H_0: $\beta_3 = 0$ and $\beta_5 = 0$ is failed to be rejected.

In all logit regressions, the bid variable itself is negative and statistically significant suggesting that the higher the dollar amount the respondent was asked to pay the less likely they would pay. This indicates a degree of internal validity of the CVM responses.

Table 3. Definition of variables

Variables	Definition of variables
Age	Age in years
CA state	Dummy variable on state of CA, 1 is CA, 0 otherwise
CA state–bid	Interaction term variable between state and bid amounts
Educ	Education of respondents (years)
ExpSmoke	Dummy variable on whether a respondent experienced smoke from forest fires or not, 1 Yes, 0 No
FL State	Dummy variable on state of Florida, 1 is Florida, 0 otherwise
FL state–bid	Interaction term variable between state and bid amounts
OwnHome	Dummy variable on whether a respondent owns home or not, 1 Yes, 0 No
RerspProb	Dummy variable determining whether a respondent has respiratory problem or not, 1 Yes 0 No
Bid	Range of bid amounts asked to pay
WitnessFire	Dummy variable on whether a respondent has seen fire or not, 1 Yes 0 No

Table 4. Logit model with pooled data of three states of CA, FL and MT for RX and mechanical programs

Variables	Coefficient (*t*-statistics)			
	Protest included		Protest excluded	
	RX	*Mechanical*	*RX*	*Mechanical*
Constant	1.3316	−0.0225	2.5524	0.1355
	(1.82)*	(−0.035)	(3.14)***	(0.21)
Age	−0.0056	−0.0006	−0.0063	−0.0018
	(−0.95)	(−0.11)	(−0.99)	(−0.35)
CA State	0.9188	0.1601	1.2129	0.1813
	(2.39)**	(0.51)	(2.75)***	(0.54)
CA Sate-Bid	−0.0011	−0.002	−0.0019	0.0016
	(−0.69)	(1.32)	(−1.08)	(1.018)
Educ	−0.025	0.0245	−0.1074	0.0274
	(−0.57)	(0.65)	(−2.21)**	(0.68)
ExpSmoke	0.0967	−0.2689	0.2589	−0.2182
	(0.35)	(−1.13)	(0.88)	(−0.89)
FL State	0.799	0.4287	0.7492	0.3063
	(2.47)**	(1.59)	(2.18)**	(1.07)
FL State-Bid	−0.0026	−0.0018	−0.0025	−0.0016
	(−1.61)	(-1.12)	(−1.43)	(−0.96)
OwnHome	0.1041	−0.1394	0.23345	−0.0952
	(0.43)	(−0.69)	(0.91)	(−0.46)
RerspProb	0.31112	0.1203	0.3938	0.1
	(1.38)	(0.63)	(1.59)	(0.53)
Bid	−0.0035	−0.0033	−0.004	−0.0036
	(−3.69)***	(−3.27)***	(−3.95)***	(−3.45)***
WitnessFire	0.0204	−0.1549	0.0601	−0.0839
	(0.092)	(−0.82)	(0.25)	(−0.43)
Mean dependent var	0.6848	0.4136	0.7229	0.4548
Log-likelihood	−369.16	−481.015	−322.59	−443.41
LR statistic (11 df)	64.3739	45.4011	74.7777	43.4057
Probability (LR stat)	1.41E-09***	4.12E-06***	1.50E-11***	9.23E-06**
McFaddenR-squared	0.08019	0.04506	0.10368	0.04666

*Significant at 10%;
**significant at 5%;
***significant at 1%.

For the mechanical program, only the bid variable is consistently statistically significant across both models. The negative sign of bid variable is as expected and indicates that the higher bid amount is asked, fewer people would pay. In term of our hypothesis test, the state variable is not significant at the 10% level in any of the regressions. The state bid interaction term is also statistically insignificant at the 10% level in all of the regressions. The geographic difference in general does not have an independent effect on support for the mechanical fuel reduction program. Thus, there may be the possibility of transferability of the WTP function for the mechanical fire fuel reduction program across the three states.

Testing equality of variable coefficients across states for the two programs

A more general test of whether the coefficients in the logit WTP models vary by state or not, is a likelihood ratio test. To do this, we estimated separate logit models for each of three states without including the state variables for two programs. The log likelihood from these models is called unrestricted ($LL_{unrestricted}$). We then run pooled data models for each program to get restricted log likelihood. Specifically, the calculated $\chi^2 = -2(LL_{restricted} - LL_{unrestricted})$.

From Table 5, it is clear that all calculated χ^2 are greater than the critical χ^2 at 1% level. Thus, there is significant difference among at least one of the coefficients across states. Therefore, the null hypothesis on equality among all the coefficients of logit models for the three states is rejected for both the RX and mechanical programs.

Mean willingness to pay comparison across states

The first step in calculation of mean WTP is to estimate separate logit models with only significant independent variables (Table 6). These models will exclude variables that have *t*-statistics less than one, as inclusion of these will unnecessarily inflate the variance and confidence intervals. As is commonly done, the mean WTPs are

Table 5. Likelihood ratio test of coefficient equality across state for RX and mechanical programs

Models	All 3 models	CA vs. FL	CA vs. MT	FL vs. MT
RX program with including protest responses				
Calculated χ^2	71.84	42.7	40.92	53.88
Mechanical program with including protest responses				
Calculated χ^2	62.72	39.5	33.9	46.82
RX program without including protest responses				
Calculated χ^2	169.7	106	101.28	125.2
Mechanical program without including protest responses				
Calculated χ^2	138.9	84.5	84.34	103.9
Critical χ^2 at *1%*	31.99	20.1	20.1	20.1
Degree of freedom	16	8	8	8

Table 6. Results of regression with significant variables for three states

RX program			Mechanical program		
Variable	*Coefficient*	*t-statistic*	*Variable*	*Coefficient*	*t-statistic*
California					
Constant	2.258908	6.46673***	Constant	−0.05041	−0.0713
RXBid	−0.00566	−3.9891***	MechBid	−0.00256	−2.0775**
			Age	0.015921	1.44873
			Expsmoke	−0.58115	−1.5864
Florida					
Constant	4.353172	3.39821***	Constant	2.958628	2.48077**
RXBid	−0.00697	−4.7731***	MechBid	−0.00399	−2.7582***
Educ	−0.21097	−2.5685***	Educ	−0.17462	−2.1751**
Ownhome	0.726814	1.94756**	Witnessfire	−0.57889	−2.0728**
Respprob	0.848499	2.09001**			
Montana					
Constant	1.283829	5.47467***	Constant	−0.2201	−0.7234
RXBid	−0.004034	−4.0563***	MechBid	−0.0037	−3.4328***

*Significant at 10%;
**significant at 5%;
***significant at 1%.

calculated for the option of excluding protest responses for each of two fuel reduction programs (Mitchell and Carson, 1989).

The formulae (Hanemann, 1989) are used for calculating the mean WTP of two programs.

$$\text{Mean WTP} = (\ln(1 + \exp(\alpha)))/B, \tag{16}$$

where α is the product of the coefficient and mean values of all independent variables excluding the bid coefficient. B is the absolute value of the bid coefficient. By using this formula, WTP_{CA}, WTP_{FL}, WTP_{MT} for each program and each option have been calculated and results are in Table 7.

The confidence intervals of 90% were calculated using a simulation technique developed by Park et al. (1991) that uses the constant and means of independent variables from computed regression outcomes and the variance–covariance matrix.

Looking at 90% confidence intervals around the mean WTP in the three states, it is apparent that these confidence intervals overlap each another. This tells us there is no statistical difference between the mean WTPs for residents of these states despite of the differences in mean WTP. In particular, the mean WTPs of people in CA and MT for the RX program are quite similar to each other; the mean WTP of people in FL and MT are quite similar to each other for the mechanical program.

Table 7. Mean annual WTP for RX and mechanical program and 90% confidence intervals

	Mean ($)	90% Confidence intervals
RX program		
CA RX	416.95	334.88–598.94
FL RX	305.04	256.34–436.24
MT RX	382.08	297.39–553.05
Mechanical program		
CA Mechanical	402.97	260.56–1308.3
FL Mechanical	229.74	159.45–515.49
MT Mechanical	207.94	157.12–338.77

Conclusion and policy implication

In order to quantify the benefits of two forest fire fuel reduction programs and test benefit transferability, we used the dichotomous choice referendum contingent valuation technique. To analyze the WTP responses, the binary logit models have been estimated for each proposed program in California, Florida and Montana. The response rate analysis consisted of the initial interview and in-depth interview responses. For people in CA, FL and MT, the χ^2 statistic of the first wave response rate ($\chi^2 = 69.89$) is significantly different at the level of 1% and 5%. However, in the in-depth WTP interviews, the χ^2 statistic is not significant at the level of 1% and 5% meaning that response rates among people in CA, FL and MT are not significantly different for this phase.

The χ^2 of protest versus non-protest responses for each of proposed programs were calculated and compared to the critical values. For the RX and mechanical programs, there was no statistically significant difference among people in three states CA, FL and MT in the pattern of protest and non-protest reasons for refusing to pay for these programs.

The next hypothesis we evaluated was whether the state of residence had an influence on voting for two proposed programs of fire fuel reduction. The logit models with including state and bid–state interaction variables were estimated with pooled data from three states of CA, FL and MT for RX and mechanical programs. For prescribed burning, the state logit intercepts do shift up the logit functions but do not rotate these bid functions in comparison to the Montana case. This says that the geographic difference has a limited impact on probability of voting for this proposed program. For the mechanical program, state variables and bid–state interaction variables are not significant at 10% level showing us that geographic difference does not have an independent effect on support for this program.

To see if the coefficients of logit models vary with state variables or not, we performed the likelihood ratio test. All calculated χ^2 are greater than critical ones at level of 1% saying that there is significant difference among at least one coefficient and these coefficients vary with state of residence.

Mean willingness to pay has been computed for the option excluding protest responses. We found confidence intervals overlap each other in three states of CA, FL and MT for the two fire fuel reduction programs suggesting no significant difference in WTP between states for the two programs.

The question raised initially in this paper is whether the WTP values and functions are transferable among three states or not? Table 8 summarizes all of the tests of transferability. A Yes means the test suggests it is transferable. Looking at Table 8, for prescribed burning program, two criteria of intercept shifter and likelihood ratio test say that survey responses are not transferable among three states of CA, FL and MT (a t-test of equality of the CA and FL intercepts accept transferability of these two state intercepts with each other, however). The bid–state interaction terms and mean WTP test show the transferability of WTP for three states of CA, FL and MT in our study. From economic point of view, the insignificance of bid–state interaction terms and equality of mean WTP may be more important criteria for examining the transferability of WTP among three states. Therefore, it appears that mean WTP is transferable among three states in our study for prescribed burning program (especially the two western states of California and Montana).

For the mechanical fire fuel reduction program, the three criteria of intercept shifter, bid interaction terms and WTP test all indicate transferability of WTP among three states of CA, FL and MT. Thus mean WTP is transferable among states for mechanical fire fuel reduction programs.

California, Florida and Montana are located in the West coast, East coast and Northern Rocky Mountains, respectively, of the US. Among these states there exist some demographic differences, differences in forest types, and the extent of wildfires. However, we found that willingness to pay to two for the prescribed burning programs were similar in California and Montana. The WTP for the mechanical fuel

Table 8. Evaluating transferability of WTP

Indicator	Transferability?	
	Protest response included	*Protest response excluded*
Prescribed burning in CA, FL and MT		
1.Intercept shifter (state logit intercepts)	No	No
2.Bid–state interaction terms	Yes	Yes
3.Likelihood ratio test	No	No
3.WTP test	Yes	Yes
Mechanical fuel reduction in CA, FL and MT		
1.Intercept shifter (state logit intercepts)	Yes	Yes
2.Bid–state interaction terms	Yes	Yes
3.Likelihood ratio test	No	No
3.WTP test	Yes	Yes

reduction program is nearly identical in Florida and Montana. The overall results appear encouraging toward transferability of mean WTP. It would be desirable to test WTP transferability among additional states of US before generalizing these results nationwide. This matter is left for future study. However, in the interim forest, managers may be able to use these mean WTP per household and WTP functions as a first approximation for estimating the benefits of fuel reduction policies in their geographic area.

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