# MULTICULTURAL EVALUATION OF THE PERFORMANCE OF CONTINGENT VALUATION FOR FOREST FIRE PREVENTION

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*This article evaluates the ability of contingent valuation to measure the benefits received by several ethnic groups from a prescribed burning forest fire reduction program similar to President Bush's Healthy Forest Initiative. Reasons for refusing to pay higher taxes for the prescribed burning program were not statistically different between African Americans, Hispanics, and whites. Mean willingness to pay of whites was $400, whereas for African Americans it was $505, but the difference is not statistically significant. The results suggest a substantial statewide willingness to pay by whites and African Americans for forest fuel reduction projects using prescribed burning in California.* (*JEL* Q26, Q23, J15)

## I. INTRODUCTION

The summers of 2000 and 2002 were two of the worst on record for forest fires in the United States. The summer of 2002 resulted in catastrophic wildfires burning nearly 6 million acres, destroying more than 2,300 houses, and resulting in 20 firefighter deaths. Damages and restoration costs are in the hundreds of millions of dollars. In response, President Bush proposed a Healthy Forests Initiative for Wildfire Prevention on 22 August 2002. A key element in this initiative is reduction of excess brush and ground fuels that have accumulated due to past fire suppression. This reduction in fuels is to be accomplished by mechanical thinning of the forests and by controlled or prescribed burning of the forest floor.

The policy of accelerating the amount of land to be mechanically thinned or prescribed burned to 2.5 million acres a year is not without opposition. Prescribed burning can generate significant quantities of smoke, which affects visibility and creates health problems for people with respiratory conditions. Prior initiatives to increase prescribed burning in states such as Florida and Washington have often been limited by citizen opposition due to smoke and health effects. The prescribed burning program is also expensive and costs as much as $250 per acre. Thus a policy-relevant issue is whether this time there will be sufficient public support for an active prescribed burning program to occur. This article extends previous work in Florida (Loomis et al., 2002) on the performance of contingent valuation method (CVM) in representing the views of Spanish-speaking Hispanics and English-speaking residents of Florida toward prescribed burning in two directions. First, the authors add a targeted sample of a third minority group (African Americans). Second, they inquire whether there are differences in CVM responses of Hispanics asked to take the survey in English (as is commonly done) versus in Spanish,

*The authors thank the Survey Research Center at California State University—Chico for implementation of the survey. Funding for this survey was provided by the Joint Fire Sciences Program. The analysis of the data was supported by the W1133 Regional Research project of the Colorado Agricultural Experiment Station. An anonymous reviewer provided extraordinarily valuable and constructive comments that greatly improved the thoroughness of the article.

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**ABBREVIATIONS**

CVM: Contingent Valuation Method
WTP: Willingness to Pay



Contemporary Economic Policy
(ISSN 1074-3529)
Vol. 22, No. 3, July 2004, 407–419

DOI: 10.1093/cep/byh030

typically a more native language for Hispanics. This allows the authors to better isolate the effect of language (Spanish versus English) from race/culture (Hispanics, whites, and African Americans). The results provide the most comprehensive evaluation of the performance of CVM in a multiracial and multilingual society, such as California.

The economic importance of understanding racial and language differences when making public policy decisions is growing daily. Collectively, minority groups are close to becoming the majority in many states of the United States. Many of these minority groups speak languages other than English. The U.S. Bureau of Census data indicates that 32 million adults in the United States speak a language other than English in their home. Furthermore, these multiracial populations are increasing faster than the English-speaking population in many states. Census data from 1990 to 1999 showed that on average, the Hispanic population grew by 39% in the United States, with states such as Arizona, California, Florida, New Mexico, and Texas having an even more rapidly increasing Hispanic population. Many in the Hispanic population either do not speak English or are more fluent in Spanish than in English. In this study area in California, nearly one-third of the population (11 million people) are of Hispanic or Latino origin. In the populous Los Angeles area, Hispanics represent 45% of the population. Another important racial or ethnic group in California are African Americans, representing 7.5% of the California population or 2.5 million people.

The growing importance of minority groups has been formally recognized in numerous policies, including Executive Order 12898, which requires federal agencies to evaluate environmental justice of federal actions on minority populations. To carry out this evaluation, policy makers must understand whether there are any differential effects of their projects or policies on many minority cultures. Surveys are a commonly used technique to assess the potential effects of policy actions on residents (Brainbridge, 1989).

A guiding assumption of survey methodology has been that similarity between interviewers and respondents has some influence on survey responses (Reese et al., 1986), and this similarity may increase the validity of survey responses (Hurtado, 1994). However, if the validity or accuracy of answers depend on matching racial, ethnic, and language characteristics of respondents and interviewers, this makes surveying more expensive and cumbersome. Thus the effect of race on answers to general surveys has been a topic of general survey research interest for decades (Lenski and Leggett, 1960). Differences in responses to racial or sensitive survey questions have been traced to differences between the race of the respondent and that of the interviewer in in-person interviews (Schuman and Converse, 1971). However, with respect to responses to questions asking for factual data or demographics, the limited research suggests no differences due to ethnicity or language for these types of questions (Weeks and Moore, 1981). There is less testing of racial/ethnic interviewer effects in phone interviews (Cotter et al., 1982), however. Further, effects of ethnicity and language on survey response rates appear not to have been systematically researched (e.g., there is no mention of this topic in Dillman's recent [2000] book and database searches do not show any entries). The authors suspect there would be response rate differences to mail and phone surveys due to race and ethnicity. Besides any obvious language difficulties, many minority cultures often feel marginalized by the dominant public institutions, such as government agencies and universities. As such they tend to have low voter participation. It is plausible that the same disinterest may carry over to answering referendum contingent valuation surveys, particularly if sponsored by the dominant culture's institutions, like government or universities.

Performing surveys in multiple languages for multiple cultures is also a difficult task. To increase within–cultural group accuracy and survey response rates and reduce item nonresponse, the survey instrument may need to be fairly culturally specific (Tindigarukayo, 2001). Yet doing so reduces its functional equivalency across groups, making group comparisons or aggregation of results more difficult (Tindigarukayo, 2001).

Given the difficulties of making surveys culturally context-specific, maintaining sufficient consistency for aggregation, coupled with the costs of training interviewers in different languages, it is not surprising that nearly all U. S. CVM surveys have been in English. Those not able to speak, read, or write proficiently in English have either been purposely omitted

in the sample design as ineligible or the individuals themselves have simply chosen not to respond. This is not just a problem in CVM. In the recreation literature, the authors are aware of very few surveys conducted in Spanish, even in California (Chavez, 2002). Whether it is explicit omission of non–English speakers or implicitly through language selection effects, this potentially leads to unrepresentative samples that limit generalizability of empirical results or, more commonly, an underestimate of benefits by omitting benefits received by non–English-speaking households. To date, no studies have compared CVM responses of whites, African American, and Hispanic households.

One might expect differences in responses across cultures to arise at any one of several stages in a CVM analysis. First, there may be differences in survey response rates. Second, there may be differences in protest refusals to pay due to a minority culture's beliefs about the ability of government to deliver on the proposed program or the fairness of the proposed method of payment. As noted by Tindigarukayo, verbatim translation of identical words into different languages may convey different meanings to whites than Hispanics. All of these subtleties can lead to differences in interpretation of the CVM scenarios and differences in willingness to pay (WTP).

The objective of this study is to determine if differences exist in survey response rates, overall WTP question protest responses, and differences in WTP estimates for white households, African American households, and Spanish-speaking households (half of which took the survey in English, as is typical, and half in their native Spanish language). The wildfire fuel reduction program under study is quite relevant to people living in California due to the state's frequent wildfires. The unique feature of this experimental design of having half the Hispanics take the survey in English and half in Spanish will aid in understanding how a respondent's native language may shape participation and response in CVM surveys. If cultural differences are found, it may suggest the need to tailor material in the CVM survey so as to better communicate with each culture. This research will also allow us to evaluate how well traditional nonmarket valuation methods (such as CVM) work with different racial and ethnic groups. The methodological approach demonstrated here has broad applicability to other fire-prone states with large Hispanic populations, such as Arizona and New Mexico.

## II. HYPOTHESES REGARDING RESPONSE RATE AND PROTEST RESPONSES

The survey modes involve an initial random-digit-dialing phone call with a short (5 minutes) initial interview. The interviewers then request their name and address to mail a survey booklet and schedule a time for an in-depth (20-minute) interview. Thus the first basis of comparison is whether African Americans, Hispanics, and whites respond equally to the initial phone call and follow through on the in-depth interview. Because the interviewers identified themselves as being with a California university, it is hypothesized that these three groups might react differently to a request from a university. Therefore they may not be equally responsive to the request for an initial interview and a follow-up in-depth interview. Such a differential response rate would make it more difficult to generalize resulting economic values from the survey sample to the population. The null hypothesis is that the overall survey response rate ($R$) to the CVM survey is independent of language and ethnicity:

$$H_0: R_{\text{AfricanAmericans}} = R_{\text{Hispanic-Spanish}} = R_{\text{Hispanic-English}} = R_{\text{White}}.$$

This will be tested using separate $4 \times 2$ contingency tables and $\chi^2$ tests for both the first and second interviews. In testing the respondent's decision to follow through at the scheduled time and participate in the second or in-depth interview, the authors can also control for demographics collected during the first initial interview. Therefore the article uses a binary logit model, where responding to the in-depth interview equals one and not responding equals zero. This is modeled as a function of race, language, age, education, number of people in the household, and gender as well as knowledge regarding forest fires in general and prescribed burning in particular. Race or language differences will be tested with a $t$-statistic on those two individual variables.

Responses to the WTP questions elicited during the in-depth interview are the main focus of analysis. First, the authors compare

the four groups reasons for refusing to pay anything. Some refusals are valid expressions of zero WTP because they reflect lack of value for the good or low income (i.e., inability to pay). Other respondents that give a zero valuation or refuse to pay because they reject the scenario or rationale that citizens should have to pay for this program, are often termed *protest responses* (Mitchell and Carson, 1989; Halstead et al., 1992). These respondents often do not "buy into" the premise that they are responsible for paying for the solution, are unconvinced the solution will actually work, or feel government will not spend the money collected on the specific program. Here as well, cultural differences between the majority culture and a minority culture may result in systematically different responses, with higher protest responses from a more distrusting minority culture.

To determine what might potentially be a protest response, the following strategy was used in the voter referendum CVM question sequence. First, if respondents indicated they would vote against the program at their initial bid amount, they were asked whether they would pay $1. If they said they would not pay $1, they were asked an open-ended question: "Why did you vote this way?" The interviewer was instructed to type in exactly what the respondent said. After all interviews were completed, the reasons were analyzed for content to classify answers by similar reasons given by the respondent. This open-ended response approach avoids having respondents fit themselves into preset protest categories or having the interviewer do so.

Comparing the overall protest reasons given, the authors will test the null hypothesis of no difference between the four groups in terms of acceptance of the premise and credibility of the CVM survey. The null hypothesis is that the distribution of refusals to pay and protest responses to the CVM survey are independent of ethnicity and language:

$$H_0: Protest_{\text{AfricanAmericans}} = Protest_{\text{Hispanic-English}} = Protest_{\text{Hispanic-Spanish}} = Protest_{\text{White}}$$

This will be tested using a $4 \times 2$ contingency table. The significance test will be performed using a $\chi^2$ statistic.

## III. WTP MODEL AND RELATED HYPOTHESIS TESTS

As suggested by a panel on contingent valuation, a voter referendum WTP question format was used (Arrow et al., 1993). Hanemann (1984) and Cameron (1988) both provide motivations for how a respondent may answer a dichotomous-choice CVM question. Hanemann views the respondent as evaluating the difference in utility associated with the status quo versus paying some amount ($\$X$) to have the program. If the difference in utility is positive for the program, the individual would respond "yes." If the difference in utility is distributed logistically, a logit model can be used to estimate the parameters and allow for calculation of WTP. The effect of language and ethnicity will be tested for using a logit model in two primary ways. First, one can test whether ethnicity and language simply shifts the logit function up or down by an intercept shifter (e.g., $B_2$, $B_4$, $B_6$) or affects the bid slope of the logit function (e.g., $B_3$, $B_5$, $B_7$) in equation (1):

$$\begin{aligned}(1)\qquad \ln(P_i/1-P_i) &= B_0 + B_1 Bid \\ &\quad + B_2 AfricanAmericans \\ &\quad + B_3 Bid * (AfricanAmericans) \\ &\quad + B_4 Hispanic\text{-}Spanish \\ &\quad + B_5 Bid * (Hispanic\text{-}Spanish) \\ &\quad + B_6 Hispanic\text{-}English \\ &\quad + B_7 Bid * (Hispanic\text{-}English) \\ &\quad + \ldots B_n X_n + u_i\end{aligned}$$

where *Bid* is the dollar amount the respondent is asked to pay, *AfricanAmericans* is one for African Americans and zero for whites and Hispanics, *Hispanic-Spanish* is one for Hispanics taking the survey in Spanish, and *Hispanic-English* is one for Hispanics taking the survey in English. The null hypotheses are

$$H_0: B_2 = 0; \quad H_0: B_3 = 0; \quad H_0: B_4 = 0;$$
$$H_0: B_5 = 0; \quad H_0: B_6 = 0; \quad H_0: B_7 = 0.$$

The hypotheses are tested through evaluation of the *t*-statistic on the respective coefficients.

A more general test is to evaluate whether ***all*** the coefficients in the logit equation would vary with ethnicity and language. Thus, four separate logit equations are estimated, one each for whites ($W$), African Americans ($AA$), Hispanic-Spanish ($HS$), and Hispanic-English ($HE$) of the form

$$(2a)\quad \ln(P_i/1-P_i) = W_0 + W_1 Bid + W_2 X_2 + W_3 X_3 + \ldots W_n X_n + u_i,$$

$$(2b)\quad \ln(P_i/1-P_i) = AA_0 + AA_1 Bid + AA_2 X_2 + AA_3 X_3 + \ldots AA_n X_n + \varepsilon_i,$$

$$(2c)\quad \ln(P_i/1-P_i) = HS_0 + HS_1 Bid + HS_2 X_2 + HS_3 X_3 + \ldots HS_n X_n + \gamma_i,$$

$$(2d)\quad \ln(P_i/1-P_i) = HE_0 + HE_1 Bid + HE_2 X_2 + HE_3 X_3 + \ldots HE_n X_n + \phi_i.$$

The null hypothesis is of coefficient equality across all four groups:

$$\mathrm{H}_0: W_0 = AA_0 = HS_0 = HE_0;$$
$$W_1 = AA_1 = HS_1 = HE_1; \ldots$$
$$W_n = AA_n = HS_n = HE_n.$$

The null hypothesis is tested using a likelihood ratio test comparing the separate logit equations to a pooled logit equation of all four groups. The significance test is conducted by evaluating the $\chi^2$ statistic. If this null hypothesis is rejected, then it is sensible to investigate which ethnicity and language treatments are the ones that are statistically different from each other and which, if any, are not statistically different from each other. Thus the authors conduct a series of pairwise likelihood ratio tests if needed.

Comparisons of mean WTP estimates across ethnicity and language groups will be used to establish if differences exist in the benefits of the public program to each group. The null hypothesis tests whether the WTP estimate by ethnicity and language are equal:

$$\mathrm{H}_0: WTP_{\text{White}} = WTP_{\text{Hispanic-English}} = WTP_{\text{Hispanic-Spanish}} = WTP_{\text{AfricanAmericans}}.$$

The results are determined by whether the confidence intervals overlap.

## IV. SURVEY DESIGN

The survey booklet was developed in conjunction with forestry professionals in California. It described the acreage that is burned by wildfires in an average year as well as the typical number of houses lost to wildfire each year. Next, a program increasing the use of prescribed fire or controlled burning in California was described. Specifically, respondents were told that the prescribed burning fuel reduction program would reduce potential wildfire fuels through periodic controlled burning. It was acknowledged that prescribed burning does create some smoke, although far less than a wildfire. Then the survey booklet provided additional information and drawings contrasting wildfire and prescribed fire. The cost of financing this program of prescribed burning was described as a cost-share program between the state of California and the county the individual lived in.

The WTP elicitation wording was:

> California is considering using some state revenue as matching funds to help counties finance fire prevention programs. If a majority of residents vote to pay the county share of this program, the Expanded California Prescribed Burning program would be implemented in your county on federal, state, and private forest and rangelands. Funding the Program would require that all users of California's forest and rangelands pay the additional costs of this program. ... If the Program was undertaken it is expected to reduce the number of acres of wildfires from the current average of 362,000 acres each year to about 272,500 acres, for a 25% reduction. The number of houses destroyed by wildfires is expected to be reduced from an average of 30 a year to about 12. Your share of the Expanded California Prescribed Burning program would cost your household $___ a year. If the Expanded Prescribed Burning Program were on the next ballot would you vote ___In favor ____Against?

The basic format of the survey booklet and script had previously been through six focus groups in two different states (including the

**TABLE 1**
Response Rates by Ethnic Group and Language with Chi-Square Tests

| | African American | Hispanics | | Whites | Total |
|---|---|---|---|---|---|
| | | English | Spanish | | |
| *First wave—screener* | | | | | |
| Total initial sample contacted | 708 | 733 | 620 | 794 | 2855 |
| Completed initial | 308 | 421 | 468 | 328 | 1525 |
| **1st wave resp rate (%)** | **43.5** | **57.4** | **75.5** | **41.3** | **53.4** |
| **Chi-square total** | | | | | **58.61*** ** |
| **Chi-sq** African Am. vs whites | **0.298** | | | | |
| **Chi-sq** Hispanic in English vs Hispanics in Spanish | | **9.98*** ** | | | |
| *Second wave—in-depth interview* | | | | | |
| Refused to give address | 4 | 9 | 1 | 4 | 18 |
| Phone disc, moved, not avail | 25 | 37 | 47 | 16 | 125 |
| Not called by end | 3 | 0 | 0 | 51 | 54 |
| **Net sample for 2nd** | **276** | **375** | **420** | **257** | **1328** |
| Completed | 126 | 170 | 139 | 187 | 622 |
| **2nd wave resp rate (%)** | **45.7** | **45.3** | **33.1** | **72.8** | **46.8** |
| **Chi-square total** | | | | | **34.25*** ** |
| **Chi-sq** African Am. vs whites | **10.51*** ** | | | | |
| **Chi-sq** Hispanic in English vs Hispanics in Spanish | | **5.48** ** | | | |

**Significant at the 5% level.
***Significant at the 1% level.

Spanish-language version), so it was only necessary to pretest the booklet and script on the four ethnic subgroups in English and Spanish prior to beginning survey administration.

The survey was conducted through a phone-mail-phone process. To obtain a representative sample of households, random-digit dialing of the households living in a sample of California counties was performed. The counties were selected so there was a mix of counties that frequently experience wildfires, counties that occasionally experience wildfires, and counties that almost never experience wildfires. Once initial contact was established, language was verified along with elicitation of initial attitude and knowledge of wild and prescribed fire, followed by the scheduling of appointments with individuals for detailed follow-up interviews. During the interim time period, a color survey booklet was mailed to the household. These interviews were conducted with the aid of this color booklet. The booklet was sent in English to whites, African Americans, and approximately half of the Hispanic households. The other half of the Hispanic households were sent the survey booklet in Spanish. The individuals were asked to read the survey booklet prior to the phone interview. Phone interviews were conducted in either English or Spanish depending on the language of the booklet received.

## V. RESULTS

### *A. Comparison of Survey Response Rates*

Because the survey was conducted in two waves, the authors compare the ethnic groups on response rates from the initial random digit dial phone survey and the follow-up in-depth interviews separately in Table 1. Although the response rates to the initial phone calls were all over 40%, there was a statistically significant difference between the four groups in response to the initial phone call. The highest response rate (75.5%) was by Hispanics phoned by a Spanish-speaking interviewer. Thus the additional effort to contact people in their native language certainly paid off during the initial interview. In particular, the study obtained a significantly higher response rate (based on the $\chi^2$ statistic of 9.98) by conducting the initial interview with Hispanics in Spanish as compared to English. Unfortunately, the

**TABLE 2**
Logit Analysis of Decision to Respond to In-Depth Follow-Up Survey

| Variable | Coefficient | *t*-Statistic | Probability |
|---|---|---|---|
| Constant | −1.2219 | −3.907 | 0.0001 |
| Age | 0.0205 | 5.189 | 0.0000 |
| African American | −0.2160 | −1.072 | 0.2835 |
| Education | 0.0341 | 1.633 | 0.1023 |
| Heard about prescribed burning | 0.3320 | 2.462 | 0.0138 |
| Heard about fire in CA | −0.0492 | −0.332 | 0.7396 |
| Hispanic English | −0.1187 | −0.614 | 0.5391 |
| Hispanic Spanish | −0.1150 | −0.670 | 0.5027 |
| Number in household | 0.0067 | 0.201 | 0.8405 |
| Bothered by smoke | −0.3907 | −2.844 | 0.0044 |
| Gender | 0.0512 | 0.453 | 0.6505 |
| Mean dependent var | 0.4077 | Log likelihood | −905.21 |
| SE of regression | 0.4793 | Restricted log likelihood | −945.07 |
| McFadden $R^2$ | 0.0422 | LR statistic | 79.720 |
| Observations with Dep = 0 | 828 | Probability | 5.69E-13 |
| Observations with Dep = 1 | 570 | | |
| Sample size | 1398 | | |

opposite effect occurs in the in-depth interviews. On mailing a Spanish language booklet to Hispanic households, the study obtained a relatively low response rate of 33% in this phase, significantly lower than obtained from Hispanics who were sent the survey booklet in English. This initially appeared somewhat counterintuitive because one would have expected that the prospect of being called back by a Spanish-language interviewer would have resulted in a higher response rate on the in-depth interview phase. However, the use of Spanish language on the initial call resulted in such a high response rate (75%) that the authors may have recruited some people who typically would not answer surveys. When they received the survey booklet and saw it was substantial (seven pages), they declined to participate in the in-depth interview. Whereas in the initial calls to Hispanics in English, the study obtained a lower initial response rate (57%), but it may have been people who were serious about completing the entire survey process. On the flip side, the response rate of whites was the highest in the in-depth interviews. Overall, the $\chi^2$ statistic suggests there is a statistically significant difference in response rates to the initial phone interview across the four samples ($\chi^2 = 58.61$) and the in-depth phone interview ($\chi^2$ 34.25). However, in the in-depth interviews, whites have a statistically higher response rate than African Americans, whereas in the initial interviews they were not different.

To investigate whether these differences in response rates persist when controlling for demographics, a binary logit model was run for the decision to respond to the follow-up in-depth CVM interviews. As indicated in Table 1, a sizable portion of the sample contacted initially did not make themselves available at their scheduled time or refused to answer the phone at that time and ignored repeated attempts to contact them at subsequent times. The logit model, where the dependent variable is one for those that completed the in-depth interview and zero for those that did not, was run including ethnicity, language, demographics, and forest knowledge/interest. As shown in Table 2, neither race nor whether the language was Spanish or English for Hispanics, has a statistically significant effect on the decision on whether to respond to the in-depth CVM interview. Thus response rate differences that show up in the lower part of Table 1 disappear when controlling for demographics. Running a logit model with just race variables (African Americans, Hispanic), and just race and language variables (African Americans, Hispanics-English, Hispanics-Spanish) by themselves indicated that the race and language variables had a negative and statistically significant ($p < 0.01$) effect on the response rate (results available from

the first author), consistent with the lower portion of Table 1. Thus it appears that demographic factors such as higher education and being older have more influence on the decision to complete the in-depth CVM interview than race and language. Higher education levels may provide some sympathy for the usefulness of social science research in the policy-making process, and being older may suggest lower opportunity costs of time.

In the Florida study, Loomis et al. (2002) found no statistical difference in response rates to either survey wave, using both the univariate $\chi^2$ test and a multivariate logit. The overall higher response rate among Hispanics and whites in Florida may be due to the older age of Florida residents and higher proportion being retirees (35% in Florida sample versus 21% in California sample), suggesting a lower opportunity cost of time in Florida. Also there was a lower average income of Florida households relative to California households, further suggesting lower opportunity cost of time might be a factor in the higher Florida response rates. The end of the article discusses why there may be other differences between California Hispanics and Florida Hispanics in terms of response rates and answers the WTP questions.

### *B. Reasons Why Households Would Not Pay for the Program*

Table 3 presents the analysis of refusals to pay, that is, individuals that indicated they were in favor of the prescribed burning program at no cost but then would neither pay their initial bid amount nor pay $1 in the follow-up WTP question. These individuals appear to favor the program but essentially have a zero WTP. Table 3 lists the reasons why a person would not pay the $1. The first three reasons listed are ***not*** considered protest responses because having no value for the program or receiving no benefits from the program, as well as not being able to afford to pay, are valid reasons for zero WTP. However, the other three categories of responses (italicized in Table 3) are considered protests because they were frequently prefaced with "I am in favor of program" or "I'm all for it, but I think the program should be paid for by those living in the forests or with existing taxes."

Because of the frequency of zero cell entries for some ethnic groups for specific protest responses, only an overall $\chi^2$ of protest responses versus nonprotest responses can be computed. The calculated $\chi^2$ of 1.994 indicates no statistically significant difference among the ethnic groups in the pattern of protest and nonprotest reasons for refusing to pay. Nonetheless, it is noteworthy that no refusals to pay were received from Hispanic households being interviewed in Spanish. Across all four groups there is substantial overall support for prescribed burning as a means to reduce wildfire. Specifically, there were only 23 households out of 622 (3.7%) that would not pay $1, and only 12 of these were considered protest responses. There were 116 people who said no to their initial bid, but of these, 92 said yes to the $1.

### *C. Results of Logit Regressions*

Table 4 presents the results of the "full" logit model that includes not only the ethnicity and

**TABLE 3**
Why Respondents Would Not Pay $1

| Reason | African Americans | Hispanics in English | Hispanics in Spanish | Whites | Total |
|---|---|---|---|---|---|
| No value/no benefits | 0 | 3 | 0 | 1 | 4 |
| Cannot afford | 1 | 1 | 0 | 3 | 5 |
| Taxes already too high | 0 | 0 | 0 | 2 | 2 |
| *Should be paid for with existing taxes* | *2* | *0* | *0* | *4* | *6* |
| *Those that live in forest should pay* | *1* | *0* | *0* | *0* | *1* |
| *Other* | *1* | *3* | *0* | *1* | *5* |
| Total | 5 | 7 | 0 | 11 | 23 |

*Note: Italicized considered protest responses for purposes of the $\chi^2$ analysis.*

**TABLE 4**
Logit Function with Ethnicity Intercept and Bid Slope Interactions
(Dependent Variable: Prescribed Burn Vote)

| Variable | Coefficient | *t*-Statistic | Probability |
|---|---|---|---|
| Constant | 2.5123 | 3.030 | 0.0024 |
| *Bid* | −0.0052 | −3.523 | 0.0004 |
| African American (*AA*) | −0.0489 | −0.089 | 0.9288 |
| *AA Bid* | 0.0010 | 0.464 | 0.6425 |
| Hispanic-English (*HE*) | −0.0995 | −0.186 | 0.8521 |
| *HE Bid* | 0.0029 | 1.387 | 0.1654 |
| Hispanic-Spanish (*HS*) | 0.4861 | 0.717 | 0.4730 |
| *HS Bid* | 0.0029 | 1.154 | 0.2484 |
| *Gender* | −0.3114 | −1.211 | 0.2256 |
| *Home Value* | 4.17E−07 | 0.767 | 0.4426 |
| *Health Problems from Prescribed Burning* | −0.8459 | −2.572 | 0.0101 |
| *Witnessed Fire* | 0.1649 | 0.585 | 0.5583 |
| *Neighbor House Burn* | 0.5292 | 1.057 | 0.2904 |
| *Income* | −5.08E−06 | −1.383 | 0.1666 |
| *Age* | −0.0022 | −0.245 | 0.8060 |
| *Number in Household* | 0.1115 | 1.262 | 0.2069 |
| *Education* | −0.0982 | −2.033 | 0.0420 |
| *FireTown* | 0.2375 | 0.690 | 0.4899 |
| Mean dependent var | 0.79 | Log likelihood | −214.57 |
| SE of regression | 0.386 | Restricted log likelihood | −237.46 |
| McFadden $R^2$ | 0.096 | LR statistic (17 df) | 45.787 |
| Sample size | 474 | Probability (LR stat) | 0.0001 |

language variables but also other demographic variables (*Age*, *Education*, *Home Value*, *Gender*, *Income*, and number of people in the household), attitudes (prescribed burning causes health problems—*Health Problems from Prescribed Burning*), as well as whether they have witnessed a forest fire (*Witnessed Fire*) and observed their neighbor's house burning (*NeighborHouseBurn*). Finally, the authors also included a variable (*FireTown*) that reflects whether the household lives in a town that has either had a fire in the recent past (e.g., Oakland) or lives adjacent to forests that repeatedly burn. These other nonethnicity variables were included to attempt to control for as many of these factors as possible to guard against our hypothesis tests of ethnicity and language being influenced by omitted variable bias.

Overall, the coefficient on the bid amount (*Bid*) is negative and statistically significant at the 0.01 level. Education level and whether respondents view prescribed burning to cause health problems from smoke were both significant at the 0.05 level. In terms of the hypotheses regarding ethnicity and language, Table 4 indicates that none of the ethnicity or language logit intercept shift variables or logit bid slope interaction terms are statistically significant at conventional levels. Specifically, the Hispanic-Spanish (*HS*) and Hispanic-English (*HE*) intercept shifters and bid slope interaction terms (*HSBid* and *HEBid*) were not statistically different from zero. Neither the African American intercept shifter nor bid slope interaction (*AABid*) variables were statistically different from zero.

The results in Table 4 suggest that if there are any differences in ethnicity and language, they may not be adequately accounted for solely by a simple intercept shifter and bid slope interaction terms. Thus the differences might be more pervasive, involving differences in all the coefficients. Using a specification focusing on just the variables in Table 4 with *t*-statistics greater than one, the authors estimated a reduced model for each of the four groups individually without, of course, the ethnicity and language variables. These logit model results are reported in Table 5.

**TABLE 5**
Logit Regressions Used for Likelihood Ratio Tests and to Calculate WTP

| | Whites | | | African Americans | | | Hispanics-English | | | Hispanics-Spanish | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| **Variable** | **Coef.** | ***t*-stats** | **Prob.** | **Coef.** | ***t*-Stats** | **Prob.** | **Coef.** | ***t*-Stats** | **Prob.** | **Coef.** | ***t*-Stats** | **Prob.** |
| Constant | 1.9659 | 1.585 | 0.112 | 3.1756 | 2.050 | 0.040 | 1.991 | 1.557 | 0.119 | 3.0149 | 1.423 | 0.15 |
| *Bid* | −0.0055 | −3.667 | 0.000 | −0.0044 | −2.532 | 0.011 | −0.0013 | −0.831 | 0.405 | −0.0026 | −1.045 | 0.29 |
| *RX Health Problem* | −0.7147 | −0.813 | 0.416 | −0.1591 | −0.2425 | 0.808 | −0.3646 | −0.568 | 0.569 | −3.0182 | −2.798 | 0.00 |
| *Income* | 0.000006 | 1.354 | 0.175 | −0.00001 | −1.896 | 0.057 | −0.00001 | −1.721 | 0.085 | −0.00007 | −0.426 | 0.66 |
| *Gender* | −0.844 | −1.899 | 0.057 | −0.0533 | −0.0912 | 0.927 | 0.1561 | 0.3361 | 0.736 | 0.4811 | 0.731 | 0.46 |
| *Age* | 0.00889 | 0.6249 | 0.532 | −0.0014 | −0.073 | 0.941 | −0.00416 | −0.231 | 0.817 | 0.02478 | 0.883 | 0.37 |
| *# in Household* | −0.05845 | −0.3460 | 0.729 | 0.3321 | 1.39 | 0.1635 | 0.2830 | 1.711 | 0.0870 | 0.12378 | 0.567 | 0.57 |
| *Education* | −0.02990 | −0.3526 | 0.724 | −0.2079 | −2.048 | 0.0405 | −0.1136 | −1.229 | 0.218 | 0.01009 | 0.078 | 0.93 |
| Mean | 0.7482 | | | 0.767 | | | 0.8142 | | | 0.8738 | | |
| Dependent log likelihood | −70.91 | | | −46.35 | | | −62.19 | | | −33.48 | | |
| Restr. log likelihood | −80.68 | | | −53.66 | | | −67.19 | | | −42.06 | | |
| LR statistic | 19.55 | | | 14.61 | | | 10.04 | | | 17.15 | | |
| Probability (LR stat) | 0.006 | | | 0.041 | | | 0.188 | | | 0.0164 | | |
| McFadden $R^2$ | 0.1211 | | | 0.1362 | | | 0.0744 | | | 0.203 | | |
| Sample size | 143 | | | 99 | | | 140 | | | 111 | | |

As can be seen in Table 5, the bid slope coefficients are statistically different from zero at conventional levels (0.01) for whites and African Americans but not for either Hispanic group. It is encouraging that the sign is negative on bid in both Hispanic regressions, but it is unusual for the bid coefficient to be insignificant in dichotomous choice CVM responses. This is contrary to what Loomis et al. (2002) found for Hispanics living in Florida for a prescribed burning program there. One difference may be that Hispanics in Florida are predominantly from the Caribbean area, whereas Hispanics in California are from Mexico and Central America. Therefore future research should investigate more subtle cultural differences, perhaps related to country of origin, to determine whether a dichotomous choice CVM survey will work for Hispanics predominantly from Mexico and Central America. Although the significance levels are fairly low on the variables individually, the likelihood ratio statistic indicates the overall equations are significantly different than the null model at the 0.05 level for the white, African American, and Hispanics in Spanish. The separate logit model on Hispanics taking the survey in English performed poorly with a probability of the likelihood ratio statistic of 0.188, perhaps suggesting this combination of English language for Hispanics does not perform well.

To test whether the coefficients are different between all the four separate regressions and the two theoretically consistent models (i.e., the two with a statistically significant negative sign on bid) likelihood ratio tests were conducted and are reported in Table 6. Given the large number of coefficients and the high variance of each model, there appears to be no statistical difference in the sets of coefficients between the four models (calculated $\chi^2 = 33.34$ while the critical was 36.15 at the 0.05 level). This lack of statistical difference in sets of coefficients is also evident comparing just the individual whites and African American models (calculated $\chi^2$ of 11.24 and critical $\chi^2$ of 15.5). These results are consistent with what Loomis et al. (2002) found in Florida when comparing English and Spanish samples logit models as well.

### D. *Comparison of Mean WTP*

Mean WTP is computed in Table 7 for the two groups whose bid coefficient in the Table 5 separate logit regressions were statistically significant (e.g., African Americans and whites). The confidence intervals were calculated using the approach of Park et al. (1991) that involves an adaptation of the Krinsky-Robb method to dichotomous choice CVM. Because the 90% confidence interval of whites overlap the mean WTP of African Americans and vice versa, there is no statistical difference between the mean WTP by whites of $399 per year and African Americans at $505 per year. Interestingly, using the African American coefficients with demographics of whites cuts the difference in WTP in half, suggesting that demographics differences play a large role in the difference in WTP. Nonetheless, the mean WTP is a sizable amount per white and African American household, suggesting that a prescribed burning program to effectuate a 25% reduction in acres of forests burned by wildfire and a 50% reduction in the number of houses destroyed is quite valuable to these California households.

**TABLE 6**
Likelihood Ratio Tests of Coefficient Equality Across Ethnic Groups

| Groups | Log Likelihood | All 4 Groups | White vs AA |
|---|---|---|---|
| White | −70.91 | | |
| African Americans | −46.35 | | |
| Hispanic-English | −62.19 | | |
| Hispanic-Spanish | −33.48 | | |
| Sum of unrestricted | | −212.93 | −117.26 |
| Pooled-restricted | | −229.61 | −122.88 |
| Calculated $\chi^2$ | | **33.34** | **11.24** |
| Critical $\chi^2$ @ 5% | | 36.15 | 15.5 |
| Significantly different? | | No | No |

**TABLE 7**
Mean WTP and 90% Confidence Intervals for Prescribed Burning Program in California

| | Mean | 90% Confidence Interval |
|---|---|---|
| Whites | $399 | $312–$608 |
| African Americans | $505 | $363–$1,126 |

## VI. CONCLUSIONS

This article investigated whether the contingent valuation method could effectively be used

to evaluate the economic effects of forest fire management policies on different ethnic groups in California. Using a univariate test ($\chi^2$), the authors found a statistical difference in survey response rates between African Americans, whites, and Hispanics for both the initial random-digit dialed interviews and the scheduled follow-up CVM phone interviews. However a logit analysis of the decisions to respond to the follow-up CVM interviews showed that once the authors controlled for demographics, response rate effects attributable to ethnicity and language disappeared. Using Hispanics' native language (Spanish) did improve the response rate to the initial interviews, but this did not carry over to the follow-up CVM interviews involving the survey booklet.

Reasons for not being willing to pay even $1 for the prescribed burning program were similar among African Americans, whites, and Hispanics taking the survey in English and Spanish. A logit regression that pooled all respondents and simply controlled for ethnicity and language using an intercept shifter variable and bid slope interaction term did not indicate any statistical differences between the four groups. A likelihood ratio test on the four separate ethnicity and language logit models confirmed that the coefficients in the logit equations are not statistically different. However, Hispanics had an insignificant coefficient on the bid, a finding robust to whether the survey was administered in English or Spanish. In contrast, whites and African Americans had a negative and statistically significant coefficient on bid amount. This may suggest that Hispanics may not have been taking the dollar amount they were asked to pay as seriously as whites and African Americans, as a significant negative sign on the bid amount is one of the few regularities in published dichotomous choice CVM surveys. This finding of insignificance on the bid amount for Hispanics in California contrasts with the finding of significance on the bid amount among Hispanics in Florida (Loomis et al., 2002). Possibly the differences in findings relates to Florida Hispanics tending to be of Caribbean descent, whereas California Hispanics are more frequently of Mexican and Central American descent. This is an interesting area of future research to determine if these cultural distinctions are the source of the difference in these findings between the two studies.

Overall, the dichotomous-choice CVM appears well suited for evaluating economic effects on African Americans and whites in California, but perhaps less well suited for evaluating the economic effects on Hispanics in California. Of course this conclusion would benefit from replication with different public policies to see if these results are specific to forest fire prevention programs or robust to the type of public good offered.

With regard to President Bush's Healthy Forest Initiative, forest fire prevention using prescribed burning does appear to have substantial support across all four sampled ethnic groups in California. Only 4% of the sample would not pay at least $1 for the program. The mean WTP of whites and African Americans is substantial at $399 and $505, respectively. Even if nonrespondents have zero willingness to pay, $400 per household over 5 million paying households would be sufficient to cover the costs of the expansion in prescribed burning in California.

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