Gasoline Prices and Purchases of New Automobiles*

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I. Introduction

The primary purpose of this paper is to measure the impact of the price of gasoline on the mix of new automobile purchases. The method of conditional logit estimation is used to model the choices of individual households between four, six, and eight cylinder cars. Given the size of the automobile industry and its influence on the rest of the economy, the determinants of the mix of automobile demand are of interest in themselves. They are especially important currently, because changes in the makeup of the automobile stock will directly affect the rate of gasoline consumption by American households.

Most recent studies of automobile demand have employed quarterly or annual national time series data.[1] In contrast, the data base used here consists of 1257 new car purchases ranging in time from January 1971 to March 1974; it thus has both cross-section and time series aspects. Although the observations span a portion of the oil embargo period of late 1973 and early 1974, our emphasis is on the response of automobile purchases to gasoline price in a relatively stable or "non-crisis" atmosphere. The disaggregated nature of the data also makes it possible to examine the demand impacts of many household-level economic and demographic factors. Income, household automobile stock, education, age, and family composition are among the variables considered in addition to gasoline price.

II. The Logit Model of New Car Purchases

Our statistical analysis of auto purchase decisions is based on the conditional logit choice model. This approach has been applied to a variety of qualitative choice problems; a deriva-

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1. For one example of a micro-level automobile demand study, see Johnson [5]. In the only recent published analysis of the mix of automobile demands, Carlson [4] applies the seemingly unrelated regressions approach to aggregate quarterly time series of five car size classes for the period 1965-1975. However, Carlson's results concerning the role of fuel prices are difficult to interpret. In his model luxury car demand is specified as a function of the gasoline price adjusted to constant dollars, while the demands for subcompact, compact, intermediate, and full sized cars are functions of the undeflated gasoline price.

tion of the model and a discussion of the method of estimation can be found in McFadden [6]. Its use in this study is justified in the following way.

Consider a sample of individual choices among $J + 1$ purchase alternatives: buying one of J types of new cars, or not buying. We assume that the utility derived from choosing the jth alternative is given by

$$V_{mj} = a_j X_m + u_{mj} \tag{1}$$

where X_m is an N-element vector of characteristics relating to the mth observation—for example, individual income level and current gasoline price—a_j is a row vector of unknown parameters, and u_{mj} is a random variable following the Weibull probability distribution. If we assume that the chosen alternative is the one offering the highest utility level, and if the u_{mj} are assumed to be independent, the $J + 1$ choice probabilities Q_{mj} take the multinomial logit form

$$Q_{mj} = e^{a_j X_m} \Big/ \left(\sum_{k=1}^{J+1} e^{a_k X_m} \right) \quad j = 1, \cdots, J + 1 . \tag{2}$$

It is readily seen that these probabilities sum to unity for each observation. Now, under the foregoing assumptions P_{mj}, which we define as the conditional probability of purchasing a type j car, given that a car is purchased, is also multinomial logit:

$$P_{mj} = e^{a_j X_m} \Big/ \left(\sum_{k=1}^{J} e^{a_k X_m} \right) \quad j = 1, \cdots, J . \tag{3}$$

The equations (3) form the basis for our empirical work. After normalizing a_1 to be a vector of zeros, we can use a sample of M observed purchases to estimate the $(J - 1) \times N$ elements of a_2 to a_J. Define $\gamma_{mj} = 1$ if alternative (automobile type) j is the alternative chosen in the mth observation, and $\gamma_{mj} = 0$ otherwise; the log-likelihood of the sample is then written

$$\log L = \sum_{m=1}^{M} \sum_{j=1}^{J} \gamma_{mj} \log P_{mj} \tag{4}$$

and consistent parameter estimates can be obtained by maximizing (4) using the Newton-Raphson technique.[2]

Differentiation of (3) shows that a_j^r, the rth element of vector a_j, measures the impact of the rth independent variable on the relative attractiveness of the jth alternative:

$$a_j^r = \frac{\partial}{\partial X_m^r} [\log(P_{mj}/P_{m1})]. \tag{5}$$

For the purposes of this study it is assumed that $J = 3$, and that the three purchase alternatives are four, six, and eight cylinder automobiles. The specification is used because of the clear relationship between gasoline mileage and number of cylinders. In 1975, the earliest model year for which data are available in the Environmental Protection Agency's *Gas Mileage Guide* [10], cylinder class explained approximately 80 percent of the variation in miles per gallon among the 315 domestic and foreign models rated.[3] Average mileage ratings by cylinder class are shown in Table I. The figures in the table indicate that changes in the cyl-

2. Thanks go to Joseph R. Antos for providing the logit estimation package used here.

3. The exact percentages of variance explained were 79.8 percent for "city" driving, and 81.8 percent for "highway" driving.

Table I. Mileage Ratings by Cylinder Class: 1975 Model Year

NUMBER OF CYLINDERS	AVERAGE MPG (STANDARD DEVIATION) CITY	HIGHWAY	TOTAL MODELS
4	20.9 (2.7)	30.4 (4.2)	68
6	16.2 (2.2)	22.0 (2.5)	57
8	11.6 (1.5)	16.6 (1.9)	190
ALL MODELS	14.4 (1.9)	20.5 (2.6)	315

Source: [10]

inder mix of new car purchases will directly affect the fuel efficiency of the total automobile stock. Conversely, purchase of the more efficient "fours" and "sixes" should be a logical consumer reaction to increased gasoline prices.[4]

The approach described in this section has a narrow focus, in that it deals only with relative purchase probabilities and makes no attempt to explain or forecast total purchases of small or large cars. However, we retain the ability to determine whether the price of gasoline has a noticeable impact on the mix of car types in the stock of automobiles. We also are able to analyze decision making at the individual level while ignoring such factors as seasonality, which may be key determinants of the buy-no buy decision but are less likely to affect the mix of purchases.

III. Data and Hypotheses

The data base used in this study is drawn from the 1972-73 Survey of Consumer Expenditures (CES) conducted by the Bureau of the Census under the auspices of the Bureau of Labor Statistics (BLS). The CES provides expenditure information along with demographic and socioeconomic data on a two-cohort sample totalling approximately 20,000 households, each of which was interviewed several times over a period of one year.[5] A de-

4. Perhaps due to the growing significance of fuel efficiency in the automobile industry, the importance of number of cylinders as a predictor of gasoline mileage seems to have declined since the time of the CES survey. For 409 models rated in the 1979 *Gas Mileage Guide* [11], cylinder class explained only 55.2 of the variance in "estimated" MPG. However, it is noteworthy that cylinder class remains a better predictor than car size class (i.e., minicompact, subcompact, small station wagon, etc.), which explains only 31.5 percent of the 1979 MPG variance.

5. The design and contents of the CES are discussed in [3].

tailed description was obtained of all cars owned by the household at the beginning of its survey year, as well as on any cars purchased between the household's first and last interviews. From the total sample of 31,155 cars owned we confine our attention to cars which were purchased new during the survey year or the previous year. We only consider families in the 23 large metropolitan areas for which we have gasoline price data,[6] and a small number of purchases are excluded due to missing information on key variables such as date of purchase or number of cylinders. The resulting sample consists of 1257 new cars purchased between January 1971 and March 1974.

An individual purchasing an automobile is investing in a durable good. With each model of automobile is associated a flow of services (a composite of transportation, comfort, etc.) and a user cost.[7] Our central behavioral hypothesis is that in areas and time periods where the gasoline price is relatively high, the relative probability of small car purchase will also be high. This should occur for two reasons. First, when the price of fuel rises the increase in automobile user cost will be smaller for more fuel efficient models, at least in absolute terms.[8] Second, unless automobile services are a Giffen good the increased user costs of all models should lead individuals to lower the quantity of automobile services which they consume. One way they can do this is by shifting their purchases toward small cars, which can be viewed as providing a smaller service flow at a lower user cost.

For each month and metropolitan area in our data base, estimates of the average retail prices of premium and regular gasoline were derived from published and unpublished BLS data. The premium and regular prices were then combined into a single gasoline price index using fixed national weights from the Consumer Price Index (CPI). A general price index was also developed, by combining the interarea index presented in Sherwood [7] with the intertemporal CPI's published by the BLS for each metropolitan area. The ratio of the gasoline index to the general price index is then used as a month- and city-specific measure of the "real" gasoline price. In the conditional logit estimation reported below, *LRPGAS*, the gasoline price measure applicable to each automobile purchase, is measured as the logarithm of the unweighted average of the real local gasoline prices for the month of purchase and the two previous months.[9]

The other economic variable used in our analysis is income. It is hypothesized that households with larger real incomes will purchase larger automobiles, *ceteris paribus*. The income measure applied to each auto purchase observation is the logarithm of total real household income before tax, in the year of purchase.[10] As with gasoline price, the income figure

6. The areas included are: New York City, Los Angeles, Chicago, Detroit, Philadelphia, Boston, Pittsburgh, Cleveland, Washington, D.C., Baltimore, St. Louis, San Francisco, Houston, Minneapolis, Buffalo, Dallas, Milwaukee, San Diego, Seattle, Atlanta, Cincinnati, Kansas City, and Honolulu.

7. User cost theory has been applied in an analysis of aggregate automobile demand by Wykoff [12]. Individuals may differ in their evaluations of the user cost and service flow for each car model. However, this does not change the sign of the hypothesized effect of fuel price changes.

8. This analysis assumes, of course, that intertemporal and cross-sectional differences in fuel prices are not completely incorporated in purchase prices, leaving relative user costs unaffected.

9. This definition was motivated in part by the fact that until late 1973 the BLS priced gasoline only quarterly for 16 of our 23 sample cities. Also, in some cities, notably Los Angeles and Detroit, gasoline prices fluctuated widely from month to month. It was expected that under such conditions consumers would base their expectations of automobile user costs on some average of current and past fuel prices. Two alternative measures of gasoline price were also tested. The first used only the current month's price; the second used an average of the prices in the current month and the five previous months, with declining arithmetic weights. Although both these alternatives were highly significant statistically, neither was as powerful an explanatory variables as *LRPGAS*. The associated coefficient estimates were of similar magnitudes regardless of the specific price variable used.

10. Since before-tax income was the only household income variable reported in the CES for the year prior to the household's survey year, analysis of auto purchases from that period necessitated use of the before-tax measure.

from the CES was deflated by the general price index described above to obtain a real income measure.

It can be hypothesized that a household's choice between car types will be affected by whether the car is to be used as a "primary" or "secondary" vehicle. For example, a small car may be more feasible for a two-car family than for a one-car family with similar economic and demographic characteristics. Therefore, a variable *CAR*2 was computed which represents the household's total stock of cars. For a given purchase, *CAR*2 equals unity if the household owned more than one car at the end of the calendar year during which the purchase took place.

A list of all variables, with their sample means and standard deviations, is found in Table II. In addition to the variables discussed above, a number of continuous and dummy variables are included to represent household characteristics such as sex, race, age, education, number of earners, family size and composition, and location.

Table II. Sample Means and Standard Deviations

VARIABLE	DEFINITION	MEAN	STANDARD DEVIATION
Choice Alternatives			
P4	1 if 4-cylinder car	.208	.406
P6	1 if 6-cylinder car	.182	.386
P8	1 if 8-cylinder car	.609	.488
Independent Variables			
LRPGAS	Logarithm of gasoline price	-.027	.065
LRINC	Logarithm of income before tax	9.772	.599
CAR2	1 if household owns 2 or more cars	.616	.487
RFAMSZ	Average family size during survey year	3.299	1.617
FEMALE	1 if female household head	.084	.278
NONWH	1 if nonwhite household head	.126	.332
CNTCTY	1 if central city household	.310	.463
RURAL	1 if rural household	.084	.277
AGEHD	Age of household head	44.187	13.086
CLGDEG	1 if household head has college degree	.268	.443
EARN2	1 if 2 or more earners in household	.625	.484
EARN3	1 if 3 or more earners in household	.220	.415
CHFAM	1 if household includes husband and/or wife, plus own children	.586	.493
HWFAM	1 if family includes husband and wife, no children	.277	.448
NCENT	1 if in North Central census region	.336	.472
SOUTH	1 if in South census region	.193	.394
WEST	1 if in West census region	.200	.400
PRE4	1 if replacing 4-cylinder car	.057	.232
PRE6	1 if replacing 6-cylinder car	.101	.301
PRE8	1 if replacing 8-cylinder car	.290	.454

Sample size for PRE4, PRE6, and PRE8 is 822 cases.
Sample size for all other variables is 1257 cases.

Although our automobile sample extends over four model years, no attempt was made to construct a measure of the relative purchase prices of fours, sixes, and eights. This omission results from the well-known difficulties inherent in measuring price changes in automobiles: the necessity of averaging over many models, estimating the size of discounts and rebates below list prices, and correcting for annual quality improvements. To the extent that relative prices change over time, some of our coefficient estimates could be affected by specification error. For this and other reasons, Section V describes several attempts to correct for possible changes in consumer behavior over time.

IV. Estimation of the Model

Table III displays the results of maximum likelihood estimation of the model described above. The three choice alternatives are purchase of a four, six, or eight cylinder car, and these are represented by probabilities *P*4, *P*6, and *P*8.[11] As discussed in Section II, with $J = 3$ we estimate two parameter vectors, so that Table III includes two coefficients pertaining to each of the 18 independent variables. The values in Column 1 measure the influence of each variable on log(*P*4/*P*8), the "log-odds" of a four as opposed to an eight cylinder purchase.

Table III. Estimated Coefficients of Base Model

	LOG(P4/P8)		LOG(P6/P8)		
	1	2	3	4	5
VARIABLE	COEFFICIENT	A.S.E.[a]	COEFFICIENT	A.S.E.[a]	CHI-SQUARE STATISTIC[b]
LRPGAS	4.277	1.245	3.534	1.273	15.26***
LRINC	-.457	.162	-.443	.159	11.90***
CAR2	.667	.190	.496	.190	15.52***
RFAMSZ	-.027	.062	-.121	.068	3.16
FEMALE	.107	.349	.916	.332	8.10**
NONWH	-.477	.247	-.363	.255	4.63*
CNTCTY	.017	.177	.323	.176	3.57
RURAL	-.189	.294	.142	.281	.90
AGEHD	-.043	.007	-.011	.006	37.70***
CLGDEG	.386	.176	.328	.186	6.14**
EARN2	.306	.196	.397	.199	5.19*
EARN3	.581	.226	.443	.229	8.16**
CHFAM	-.237	.315	.201	.334	1.32
HWFAM	-.673	.325	.040	.337	4.81*
NCENT	-.788	.210	-.388	.198	15.22***
SOUTH	-.573	.232	-.634	.242	10.31***
WEST	.702	.215	-.125	.234	10.93***
INTERCEPT	5.140	1.495	3.309	1.487	13.13***
		NUMBER OF CASES		1257	
		LOG-LIKELIHOOD		-1092.8	

[a] Asymptotic Standard Error

[b] Values are for the statistic measuring the combined significance of the two coefficients in that row. Under the null hypothesis each statistic is asymptotically chi-square with two degrees of freedom. Values marked with one, two and three asterisks exceed the 90%, 95%, and 99% points of the chi-square distribution, respectively.

11. Note that in our micro-level data, the observations on probabilities *P*4, *P*6, and *P*8 are either zeros or ones.

Similarly, Column 3 contains estimates of the derivative of log(*P*6/*P*8) with respect to each independent variable. Thus, positive coefficients indicate that increases in the variable raise the probability that a buyer will purchase a four or six rather than an eight.

Statistical significance can be inferred by comparing a coefficient to its asymptotic standard error. In addition, the chi-square statistic in Column 5 of the table measures the combined significance of the Column 1 and Column 3 coefficients. A significant chi-square value implies rejection of the null hypothesis that both coefficients are zero—i.e., that the associated variable has no effect on relative purchase probabilities.

The large number of significant chi-square values in Column 5 suggests that, taken as a group, the independent variables are important determinants of car purchase decisions. A test of the validity of the entire model can be constructed by comparing the log-likelihood value reported in Table III to the log-likelihood obtained under the constraint that all coefficients except the intercept terms are equal to zero. The resulting chi-square test statistic is 174.7 with 34 degrees of freedom.[12] We can therefore decisively reject the hypothesis that relative purchase probabilities vary only randomly across individuals.

Predicted purchase probabilities may be obtained by inserting values of the explanatory variables into an equation of the form of (3) above. For example, let us define a "representative household" as a three-person (husband, wife, and one child), two-earner, two-car, suburban family living in the North Central region, with sample mean income and facing the sample mean price of gasoline. The household head is assumed to be a white male, aged 45, with a high school education. For such a household, the predicted purchase probabilities are 14.5 percent, 18.5 percent, and 67.0 percent for fours, sixes, and eights respectively. The influence of individual variables on these probabilities is discussed below.

(a) The coefficients on *LRPGAS* are both positive as expected, and are highly significant statistically, implying that both log(*P*4/*P*8) and log(*P*6/*P*8) are positively related to *LRPGAS*. Further, the Column 1 coefficient exceeds the Column 3 coefficient in absolute value. Thus, increases in fuel prices raise the attractiveness of fours relative to sixes, and of both fours and sixes relative to eights. A gasoline price increase of one sample standard deviation is equivalent to 6.5 percent, or about three cents in 1973 prices. Such an increase is estimated to increase the probability of four cylinder purchase to 17.5 percent, with the six cylinder probability rising to 21.3 percent.[13]

(b) As hypothesized, the coefficients on LRINC, the logarithm of income, are negative and highly significant. The implication is that the income elasticity of eight cylinder demand exceeds the elasticity of demand for the other two categories. At sample means an increase of ten percent in real before-tax income is estimated to increase the probability of eight cylinder purchase by about one percent.

(c) The *CAR*2 coefficients are positive and significant, indicating that two-car households are more likely than one-car households to purchase fours and sixes. For households without a second car, the four-cylinder purchase probability falls to 8.7 percent and the six-cylinder probability to 13.1 percent. As noted earlier, in some types of households a small car may be more feasible as a "second car" than as the only automobile owned.

12. Minus twice the difference in log-likelihood values is, under the null hypothesis, asymptotically chi-square distributed, with degrees of freedom equal to the number of constrained parameters. See, for example, Theil [9, 396–7].

13. Here, as in the rest of this section, the "probability of purchase" of a given car type is to be interpreted as the probability conditional on purchase of some new car.

(d) Two dummy variables, *CNTCTY* and *RURAL*, were introduced to test the importance of location (central city, rural, or other) within a metropolitan area. However, neither variable is a significant determinant of purchase behavior in our data.

(e) Age of household head has a strong negative effect on the demand for fours relative to eights. As age increases from 25 to 65, the probability of four cylinder purchase falls from 27.7 percent to 7.0 percent, while the predicted eight cylinder percentage increases from 53.7 to 76.3. The demand for sixes as a percentage of total demand is essentially invariant with respect to age.

(f) The estimated effect of a college education is to increase the relative demand for fours and sixes.

(g) Regional effects are highly significant. The predicted probability of eight cylinder purchase ranges from 41.3 percent in the West to 67.3 percent in the South. The four-cylinder probability is almost three times as high in the West (39.7 percent) as in the North Central region (14.5 percent). Among the potential explanations for these differences are regional variations in weather and in urban commuting patterns. It is also possible that the relative purchase prices of small cars—many of which are foreign built—are lower on the West Coast.

(h) Of the seven remaining demographic variables, five—*FEMALE, NONWH, EARN2, EARN3,* and *HWFAM*—are significant at (at least) the ten percent level, judging from their chi-square values in Column 5. Some of the evident results are preferences for sixes by women, for eights by nonwhites, for fours and sixes by multi-earner households, and for sixes and eights by husband-and-wife couples.

For every explanatory variable except *HWFAM* and the insignificant *RURAL* and *CHFAM*, the coefficients in Column 1 and Column 3 are of the same sign. That is, the same factors which favor the purchase of fours relative to eights also favor sixes relative to eights. The Column 1 and Column 3 coefficients are often also of similar magnitudes, suggesting that fours and sixes may often be viewed as approximately equivalent "small car" alternatives. For this reason, the model was re-estimated under the constraint that the Column 1 and Column 3 coefficients are equal to each other for each explanatory variable (excepting the intercepts). This permits us to make a likelihood ratio test of the significance of the whole explanatory variable set in explaining choices between fours and sixes. The null hypothesis is that the ratio *P4/P6* is a constant, and that the explanatory variables only affect *P8*. The associated chi-square test statistic enables us to reject the null hypothesis at the .001 significance level. We can therefore conclude that the model of Table III is superior to a binary model in explaining automobile demands by size class.

V. Robustness of Estimates

This section reports on the results of modifying the Section IV specification in a series of ways, through the addition of new variables or deletion of subsets of observations. In each instance the objective is to measure the sensitivity of the *LRPGAS* coefficients to the precise makeup of the sample and specification of the choice equation.

One possible criticism of the Table III estimates is that the time series aspect of the data base was not adequately addressed. Much of the sample variation in *LRPGAS*, for example, arises from a steady increase in the real gasoline price during 1973 and early 1974. There-

fore, the *LRPGAS* coefficients could be biased if there were trends in consumer attitudes or relative purchase prices favoring smaller cars during the sample period.

We tested for this sort of specification error by estimating the choice equation separately on the 1972 and 1973 CES cohorts. The resulting *LRPGAS* coefficient estimates are displayed in the second and third rows of Table IV, with the pooled estimates from the preceding section repeated in Row 1. The coefficients suggest that the impact of *LRPGAS* on log (*P*4/P8) declined slightly in 1973, while the sensitivity of six cylinder demand to the gasoline price rose sharply. However, even in the case of log (*P*6/P8) the coefficients do not differ by more than the sum of their standard errors. Further, using a likelihood ratio test the null hypothesis of complete cohort equivalence (i.e., equality of the two 36-element coefficient sets) could not be rejected. The associated chi-square value was 23.0 with 36 degrees of freedom; this value would be exceeded in over 90 percent of trials under the null hypothesis.

Behavioral trends were also tested for directly, through the inclusion of a variable *MONTH*, defined as the number of months between December 1969 and the month of purchase. When added to the Table III specification, *MONTH* had no appreciable effect. Its two estimated coefficients were much smaller than their asymptotic standard errors, and the *LRPGAS* coefficients remained essentially unaffected.

Table IV. Alternative Estimates of LRPGAS Coefficients

MODEL	COEFFICIENT (Asymptotic Standard Error) LOG(P4/P8)	LOG(P6/P8)	NUMBER OF CASES	CHI-SQUARE STATISTIC[a]
1. Base Model (Table III)	4.28 (1.25)	3.53 (1.27)	1257	15.26***
2. 1972 CES Cohort	4.36 (1.88)	1.65 (1.85)	626	5.49*
3. 1973 CES Cohort	3.56 (1.81)	4.83 (1.88)	631	8.08**
4. MONTH Trend Variable Included	4.25 (1.30)	3.53 (1.33)	1257	13.92***
5. PCHANGE Price Change Variable Included	4.25 (1.29)	3.73 (1.32)	1257	14.85***
6. Energy Crisis Period Excluded	3.75 (1.40)	3.29 (1.42)	1175	9.96***
7. Previous-Car Variables Included (Table V)	4.63 (1.65)	3.16 (1.62)	822	9.29***

[a] See note b of Table III.

Similar results were obtained from a variable *PCHANGE,* computed as the percentage change in deflated gasoline price during the three months prior to the date of purchase. This variable was intended to capture any effects of changing consumer expectations resulting from recent fuel price trends. Once again, the new variable was statistically insignificant and its inclusion had little influence on the *LRPGAS* coefficients. The results of the runs using *MONTH* and *PCHANGE* are displayed in Rows 4 and 5 of Table IV.

In a fourth experiment, all purchases which occurred after October 1973 were excluded from the sample base. The intent was to purge the data of any effects of the oil embargo and energy crisis. This truncation resulted in the loss of 82 of 1257 observations, and the sample mean and standard deviation of *LRPGAS* were reduced somewhat. The estimated coefficients on *LRPGAS* were also reduced, although they remain highly significant statistically. Two explanations can be offered for the reduction. First, the elasticity of automobile purchases with respect to gasoline price could be greater at higher gasoline price levels such as those prevailing during the embargo period. Second, a higher responsiveness after October 1973 could reflect a change in consumer expectations about the trend in gasoline prices. The time span of the CES data base does not extend far enough into 1974 to permit a clear test of either of these hypotheses. For present purposes, the important result is that a strong gasoline price effect is evident even in the pre-embargo period.

The foregoing tests have indicated that the results presented in Section IV do not suffer from an inadequate treatment of trend effects. A cross-sectional specification error could also be present, if inter-area variations in purchase behavior resulting from differences in attitudes, climate, or other factors are correlated with local gasoline prices. The purpose of including the regional dummy variables *NCENT, SOUTH,* and *WEST* was to avoid this kind of bias. Spurious locational effects can be further reduced through the introduction of variables representing the size of car replaced by the new purchase. The inclusion of such variables should hold constant purchase behavior patterns associated with location. The *LRPGAS* coefficient should then indicate whether higher gasoline prices lead to changes in historical patterns.

Table V displays the results of logit estimation with the Table III variables plus the dummy variables *PRE*4, *PRE*6, and *PRE*8 indicating whether the replaced car had four, six, or eight cylinders. The CES does not provide information on the type of car replaced for purchases prior to the household's survey year. As a result, the number of usable observations is reduced to 822, ranging in time from January 1972 to March 1974. For approximately 55 percent of these purchases the data yield no evidence of a simultaneous sale. In such cases, *PRE*4, *PRE*6, and *PRE*8 all equal zero and the type of car replaced is "none or unknown."

Comparison of Tables III and V shows that neither *LRPGAS* coefficient is severely affected by inclusion of the three previous-car variables. The estimated effect of *LRPGAS* on log $(P6/P8)$ falls slightly, but the coefficient remains almost twice as large as its asymptotic standard error despite the loss of observations. Hence, Table V provides no evidence of a serious locational bias in our earlier gasoline price coefficients. In general, coefficients on the remaining explanatory variables lose magnitude and significance in Table V as compared to Table III.

The *PRE*4, *PRE*6, and *PRE*8 coefficients perform as expected, with individuals tending to purchase the same type of car as they are replacing. For example, for a purchaser who is replacing an eight cylinder car, the estimated probabilities of four, six, and eight cylinder

purchase are 8.6 percent, 13.4 percent, and 77.9 percent, respectively. These percentages shift to 40.6, 19.1, and 40.3 for a household replacing a four. The insignificant coefficients on *PRE6* imply that purchasers who are replacing sixes behave similarly to those who are not replacing another car (or whose previous car is of unknown size). This supports the natural view of a six cylinder car as an "intermediate" type.

VI. Conclusion

The foregoing results can be taken as evidence of a strong cross-elasticity of small car demand with respect to the price of gasoline. In each of the estimation runs reported in Table IV, the *LRPGAS* coefficients were large and statistically significant. For the period covered

Table V. Estimated Coefficients of Three-Alternative Model With Previous-Car Variables

	LOG(P4/P8)		LOG(P6/P8)		
	1	2	3	4	5
VARIABLE	COEFFICIENT	A.S.E.[a]	COEFFICIENT	A.S.E.[a]	CHI-SQUARE STATISTIC[b]
LRPGAS	4.634	1.649	3.163	1.621	9.29***
LRINC	-.445	.196	-.317	.193	6.11**
CAR2	.350	.240	.435	.234	4.48
RFAMSZ	-.032	.079	-.140	.084	2.78
FEMALE	-.017	.428	.932	.372	7.16**
NONWH	-.435	.319	-.153	.300	1.87
CNTCTY	-.325	.230	.115	.218	2.95
RURAL	-.233	.392	.322	.337	1.74
AGEHD	-.043	.009	-.007	.008	21.06***
CLGDEG	.482	.227	.297	.233	4.89*
EARN2	.273	.257	.421	.248	3.29
EARN3	.694	.300	.204	.287	5.35*
CHFAM	-.041	.385	.312	.380	.81
HWFAM	-.298	.395	.069	.390	.74
NCENT	-.708	.267	-.318	.247	7.38**
SOUTH	-.694	.305	-.695	.304	8.26**
WEST	.646	.281	.183	.292	5.31*
PRE4	1.085	.369	.274	.462	9.21***
PRE6	-.127	.330	.156	.301	.59
PRE8	-1.116	.272	-.736	.239	22.47***
INTERCEPT	5.202	1.778	2.154	1.765	8.62**

NUMBER OF CASES 822
LOG-LIKELIHOOD -693.5

[a] Asymptotic Standard Error.
[b] See note b of Table III.

by the CES sample, a ten percent increase in fuel price is estimated to yield an increase of just over eight percent in the proportion of small (four and six cylinder) cars purchased. Thus, adjustment of the automobile stock should be a significant component of the total response to rising gasoline prices. This stock adjustment effect, which previously has been studied only at the aggregate level,[14] is in addition to the strong short run response confirmed most recently by Archibald and Gillingham [1] using household-level data on gasoline consumption with auto stocks held constant.

Currently the federal government is intervening directly in the new car market. Beginning in 1978, the new car sales mix of each auto maker has been required to exceed minimum fuel economy standards. An implication of our analysis is that a continuation of gasoline price increases would lessen the upward pressure on large car prices which might be expected to result from the federal mandates.

Recent years have witnessed a steady growth in small car demand for fuel efficient automobiles. In addition to the gasoline price, our results suggest several other reasons for this trend. For example, we find that small cars are relatively popular among multi-car and multi-earner households. To the extent that these types of households continue to become more popular, the trend toward small cars should also continue.

14. See, for example, Sweeney [8] and Burright and Enns [2].

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