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# ADVERTISING AND INTRAINDUSTRY BRAND SHIFT IN THE U.S. BREWING INDUSTRY*

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## I. INTRODUCTION

The brewing industry has constituted a focus of interest in industrial-organization economics since World War II, due to its rapidly increasing concentration, possibly large plant-level scale economies, and marked product differentiation accompanied by relatively high firm advertising intensities. Scherer [19, p. 110] wrote concerning the dramatic postwar structural change which has characterized this industry:

> "... That in turn permitted them (Anheuser-Busch, Schlitz, and Pabst) to build additional new breweries which, by exploiting a combination of technological advances and scale-up economies, had much lower production costs per unit than those of regional brewers lacking premium image brands. With lower costs, the Big Three (joined by Coors and later Miller's) could squeeze the premium-popular price differential, enhancing their market shares even more, which in turn permitted them to advertise as heavily in absolute terms as smaller regional rivals but at appreciably lower average outlays per sales dollar. The upshot of these interacting developments was an increase in the four-firm concentration ratio in brewing from 21 in 1947 to 61 in 1976."

Although the relative effects of these factors on the rise in concentration are debated (see, e.g., Horowitz and Horowitz [9] and [10], Greer [6], Hawkins and Radcliffe [8], and Elzinga [4]), we focus on product differentiation and advertising and their effect specifically on consumer brand selection. The results of our study shed some light on the dynamics of concentration and market-share change in brewing.

Since Telser [21] modeled consumer brand change as a stochastic process and estimated brand-change transition probabilities from consumer panel data, this *specific line* of inquiry has remained relatively dormant. In this paper, we reconsider this area of research by employing a recently-proposed technique for transition-probability estimation which allows us to examine the brand-change issue more precisely.

From monthly consumer panel data for 6,000 families, Telser estimated

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repeat- and transfer-purchase probabilities for unnamed brands of four different commodities. He argued that, although micro data permitted maximum-likelihood estimates of Markov-chain transition probabilities, these probabilities also depended on relative brand prices; he employed a linear and brand-aggregated (i.e., one brand's price is expressed relative to an average of all other brands' prices) approximation of this relative-price dependency, and estimated the model

$$x_j(t) = L_{0j} + L_{1j}x_j(t-1) + L_{2j}P_j(t) + \varepsilon_j(t);$$

$x_j(t)$ is the market share of brand $j$ at time $t$, $P_j(t)$ represents the price of brand $j$ at time $t$ minus the average price of all other brands at time $t$, $\varepsilon_j(t)$ is the stochastic sampling error, and $L_{0j}$, $L_{1j}$, and $L_{2j}$ are regression coefficients. Telser's linear approximation does not guarantee the nonnegativity and row-sum constraints of the Markov model, discussed below, nor is it clear that probability estimates obtained in this manner are statistically superior to micro maximum-likelihood estimates.

It is now possible (as shown by MacRae [15]) to estimate the time-dependent effect of economic factors on brand-change probabilities (and market shares), under the basic Markov-model constraints, even when frequency data *alone* are available: i.e., temporal snapshots of market shares. We feel that MacRae's technique can be profitably employed while focusing on consumer brand change in the brewing industry. In essence, we attempt a test of Greer's hypothesis that

> "to a great extent shifts in market shares since 1953 have reflected relative intensities of advertising" (Greer [6, p. 217]).

## II. THE MODEL

In a discrete-time Markov process there are a finite number of states indexed by $i$ ($i = 1, \ldots, R$).[1] Here, each state corresponds to a brewing firm or group of firms. For each realization of the process, the "system" may be unambiguously classified as being in one of the states $i$. Define the random variable $X_t$ to be the state of the system at time $t$. Then the probability of transition from state $i$ to state $j$ at time $t$ is

$$(1) \qquad p_{ij}(t) = Pr(X_t = j \mid X_{t-1} = i).$$

[1] Following the work of Horowitz and Horowitz [9], there is no closure or "non-beer-drinking" state specified for this model. Results are suggestive only of general industry trends. It is possible, theoretically, to add a closure state to the model even if aggregate frequency data alone are available. (See Kelton [13] for an empirical application. Also, see Kelton [13] or Collins [2] for invariance arguments and tests with respect to the exact size of the specified closure state.)

If $p_{ij}(t) = p_{ij}$ for all $t$, the Markov process is said to be *stationary*. Note that

$$(2) \qquad 0 \le p_{ij}(t) \le 1 \text{ and } \sum_{j=1}^{R} p_{ij}(t) = 1, \text{ for all } i \text{ and } t.$$

Let $P(t)$ be an $R \times R$ matrix with $(i, j)$th element $p_{ij}(t)$. Hence, (2) says that each entry of $P(t)$ lies between zero and unity, and the sum of the elements in each row of $P(t)$ is unity.

If information is not available concerning the number of individual transitions between each pair of states, the usual micro maximum-likelihood estimation technique (and associated statistical inference procedures—see Anderson and Goodman [1]) cannot be employed. With a technique proposed by MacRae [15], which has not yet been applied empirically, transition probabilities may be estimated with aggregate frequency data and an embedded independent variable. *Consistent* estimates are obtained (see MacRae [15, p. 190]) which automatically satisfy the constraints (2).

The procedure involves solving an unconstrained nonlinear programming problem to obtain estimates of the nonstationary transition probabilities.[2] This problem is essentially one of nonlinear least squares, where the residuals are the deviations of observed market shares from predicted shares in the Markov model. In formulating the problem, two types of regression coefficients are introduced. The first, $\alpha_{ij1}$, reflects the tendency for consumers to shift from brand $i$ to brand $j$, regardless of the advertising differential between these brands. The second, $\alpha_{ij2}$, reflects the advertising-induced tendency to shift from brand $i$ to brand $j$. Details of the estimation procedure are given in the Appendix.

Let $A_i(t)$ be the advertising expenditure of firm $i$ at time $t$. Our basic hypothesis is that a large $A_j(t-1)$ is associated with a large $p_{ij}(t)$, for $i \neq j$.[3] Thus, for $i \neq j$, we examine $\partial p_{ij}(t)/\partial A_j(t-1)$, which is positive if and only if $\alpha_{ij2} < 0$, by (7) in the Appendix.[4]

We argue that consumers will tend to transfer purchases from those firms that do not advertise intensively to those that do. However, as Telser argued [21, p. 301], these tendencies may be either slowly or rapidly operating, depending on the strength of consumer attachment to particular firms' brands; brand shifting may occur regardless of advertising differentials—hence, the two "components" of the tendency to shift brands: a pure "stochastic component" $\alpha_{ij1}$ and the advertising coefficient $\alpha_{ij2}$.

[2] It may be shown (see Kelton [12]) that MacRae's procedure produces estimates identical to those which could be obtained by a quadratic programming procedure in the case of stationarity.

[3] Theoretically, the model could be generalized to incorporate lags of longer than one period, although Grabowski [5, p. 692] estimated a relatively high rate of depreciation of the "goodwill stock" for the brewing industry; to incorporate a threshold (i.e., discontinuous) influence of the advertising differential, or to incorporate other "independent" variables such as relative beer prices. With sufficient data, the number of states could also be expanded.

[4] If $j$ is fixed and $\alpha_{ij2} < 0$ for all $i$, then the market share of firm $j$ also grows as its advertising expenditure increases (i.e., $\partial x_j(t)/\partial A_j(t-1) > 0$). If $R = 2$, the converse is true as well; see (3).

The null hypothesis that $\alpha_{ij2} = 0$ (i.e., no advertising effect on market share and a stationary Markov process) may be tested using regression-analysis econometric techniques (see Theil [22, pp. 137–45, 312–14] for derivation of testing procedures).[5]

## III. THE DATA

The data consist of yearly shares [from 1951 ($t = 0$) to 1977 ($t = 26$)][6] of national barrelage [$x_B(t)$ and $x_M(t)$ for Anheuser-Busch and Miller, respectively] taken from the Federal Trade Commission's report *The Brewing Industry* [11, p. 22].[7] $A_B(t)$ and $A_M(t)$, or total major-media advertising expenditure by these two brewers, is also constructed from *The Brewing Industry*. $A_0(t)$, other major-media advertising expenditure in the brewing industry at time $t$, is taken as a six-firm total: Schlitz, Pabst, Falstaff, Schaefer, Stroh, and Schmidt, for which yearly advertising cost data were presented in the FTC report.

## IV. EMPIRICAL RESULTS

A recent nonlinear programming algorithm (see Nórris and Gerken [18]) was used to solve the minimization problem (5). Various questions of interest are better addressed if we consider two- and three-state models separately.

### A. *Two States*

We initially considered a two-state system, Miller and "Other." (Thus, $R = 2$, and we let state 1 be Miller, or "M" in the ensuing subscripts; "O" stands for "Other," which is state 2.)

Table I presents the estimated "repeat-purchase" and "transfer-purchase" coefficients for Miller as the "destination state"; $\hat{\alpha}_{OM2} < 0$, as suspected. Furthermore, $\hat{\alpha}_{OM2}$ is significantly negative [the calculated F-statistic with (1, 23) d.f. was 52.9, significant at the 1% level].[8] Thus, the lagged effect of Miller's advertising on its transfer-purchase probability is significantly positive. Also, it appears that the transition probabilities are not stationary [see (6)].

[5] Since the stationary Markov-chain model is simply a special case of the Markov-process model with an embedded advertising variable when all the $\alpha_{ij2}$'s are constrained to be zero, analogy to linear regression theory suggests that under the null hypothesis,

$$\frac{(SSR_R - SSR_U)/[(R-1)^2]}{SSR_U/[(T-R)(R-1)]} \sim F_{(R-1)^2,\,(T-R)(R-1)}.$$

approximately, where $SSR_R$ and $SSR_U$ are the sums of squared residuals from the stationary and nonstationary estimations, respectively.

[6] Results are not affected by considering the shorter period 1961–1977, after a "decade of stagnation" for the national brewers, as discussed in Horowitz and Horowitz [10] and Greer [6].

[7] The original FTC sources included *Advertising Age* and *Modern Brewery Age* for various years. Advertising is measured in current dollars. Estimates are unchanged if advertising expenditures are instead measured in constant dollars, and elasticity estimates are dimensionless.

[8] This conclusion is tentative subject to more definitive Monte Carlo studies to determine the exact appropriateness of the F-distribution, and analogy to linear regression theory, in this context.

TABLE I
COEFFICIENT ESTIMATES FOR THE TWO-STATE NONSTATIONARY MARKOV PROCESS

| $\alpha_{MM1}$ | $\alpha_{OM1}$ | $\alpha_{OM2}$ |
|---|---|---|
| 14.83 | −1.07 | −.70*** |

*** Significant at the 1% level of significance.

M Miller beer.

O All other beer.

As mentioned in Note 4, $\hat{\alpha}_{OM2} < 0$ implies that Miller's market share should grow with an increase in the previous year's advertising outlay. For the present two-state case, however,

$$(3) \qquad \partial x_M(t)/\partial A_M(t-1) = x_O(t-1)[\partial p_{OM}(t)/\partial A_M(t-1)],$$

so that $\hat{\alpha}_{OM2} < 0$ is also necessary for $\partial x_M(t)/\partial A_M(t-1)$ to be positive. The magnitude of the lagged effect of a percentage change in advertising expenditures on market share at a given time (in essence an advertising elasticity) can be measured by

$$(4) \qquad E(t) = [\partial x_M(t)/\partial A_M(t-1)]A_M(t-1).$$

We used (3) to evaluate $E(t)$ for $t = 1, \ldots, 26$, which ranged from essentially zero between 1962 and 1964 to .0731 in 1977, and had an average value of .0141. Thus, $E(t)$ has not increased consistently with time, but a sharp rise in $E(t)$ which occurred in the 1970s coincides with Phillip Morris's new advertising campaign for Miller.

The estimated transfer-purchase probabilities $\hat{p}_{OM}(t)$ for Miller can be calculated from Table I, (6), and the general expression for $p_{ij}(\beta, t)$ in the Appendix; Table II gives these estimated probabilities for the later years. Note that over this period, the estimated transfer-purchase probabilities exhibit clear nonstationarity and a tendency to increase. The fitted model can also be used to predict market shares; for example, using the actual market shares for Miller and "Other" in 1977 and $\hat{P}(26)$ ($t = 26$ corresponds to 1977), assuming that $\hat{P}(26) \simeq \hat{P}(27)$, the 1978 and 1979 market shares for Miller are predicted to be .193 and .215, respectively. (Independent estimates of the actual market shares for these two years were reported to be .194 and .245 in the August 4, 1980 issue of *Advertising Age*, and .193 and .215 in a Kidder, Peabody & Co., Inc. Research Department report of January 1980.)[9]

[9] Assuming *stationary* probabilities, the transfer-purchase probability estimate is .0047, which is substantially lower than the nonstationary counterparts for the later years. Starting again in 1977, the stationary model predicts 1978 and 1979 market shares for Miller to be, respectively, .158 and .162, clear underestimates. Thus, ignoring nonstationarity is certainly inappropriate in this case.

TABLE II
ESTIMATED TRANSFER-PURCHASE PROBABILITIES 1967 TO 1977

| *Year* | *Probability Estimate* |
|---|---|
| 1967 | .0009 |
| 1968 | .0028 |
| 1969 | .0029 |
| 1970 | .0064 |
| 1971 | .0059 |
| 1972 | .0104 |
| 1973 | .0038 |
| 1974 | .0040 |
| 1975 | .0118 |
| 1976 | .0296 |
| 1977 | .0428 |

## B. *Three States*

It is interesting to expand the model to consider a three-state system, with Anheuser-Busch (B), Miller (M), and "Other" (O) as the states. The estimated coefficients may be found in Table III. Note that Miller's two "transfer-purchase" advertising coefficients (the last two estimates in Table III) are certainly larger in magnitude than those for Anheuser-Busch. Three of the coefficients (all but $\hat{\alpha}_{MB2}$) are negative, as suspected. Furthermore, the null hypothesis that $\alpha_{MB2} = \alpha_{OB2} = \alpha_{BM2} = \alpha_{OM2} = 0$ (assumption of no advertising effect and stationarity) may be tested against the alternative that at least one $\alpha_{ij2} \neq 0$.

TABLE III
COEFFICIENT ESTIMATES FOR THE THREE-STATE NONSTATIONARY MARKOV PROCESS

| *Coefficient* | *Estimate* |
|---|---|
| $\alpha_{BB1}$ | 7.853 |
| $\alpha_{MB1}$ | −4.995 |
| $\alpha_{OB1}$ | −4.547 |
| $\alpha_{BM1}$ | 2.556 |
| $\alpha_{MM1}$ | 5.855 |
| $\alpha_{OM1}$ | −0.848 |
| $\alpha_{MB2}$ | 0.008 |
| $\alpha_{OB2}$ | −0.085 |
| $\alpha_{BM2}$ | −6.264 |
| $\alpha_{OM2}$ | −1.059 |

With (4, 42) d.f., the F-value is 2.616, and the null hypothesis can be rejected at the 5% level.[10] With these coefficient estimates, we can calculate, for 1977 (i.e., $t = 26$),

$$E_B(t) = \frac{\partial x_B(t)}{\partial A_B(t-1)} A_B(t-1)$$
$$= \left\{ \sum_{i=1}^{R} \frac{\partial p_{iB}(t)}{\partial A_B(t-1)} x_i(t-1) \right\} A_B(t-1) = .0010;$$
$$E_M(t) = \frac{\partial x_M(t)}{\partial A_M(t-1)} A_M(t-1)$$
$$= \left\{ \sum_{i=1}^{R} \frac{\partial p_{iM}(t)}{\partial A_M(t-1)} x_i(t-1) \right\} A_M(t-1) = .0692.$$

Advertising's effectiveness clearly differs across firms. The "elasticity" of market share with respect to own lagged advertising expenditure appears much higher for Miller than for Anheuser-Busch. If we define the transfer-purchase probability for firm $j$ at time $t$ as

$$T_j(t) = \sum_{\substack{i=1 \\ i \neq j}}^{R} p_{ij}(t),$$

then $T_B(t)$ and $T_M(t)$ differ over time as do the elasticity estimates above. $T_B(t)$ remains relatively constant over the period, and ranges between .0072 and .0090, with average value .0083, but $T_M(t)$ rises greatly, yet not consistently, over time from .0006 for 1952 to .0568 for 1977, with average value .0096.

Moreover, the nonstationary model again offers more accurate market-share predictions for 1978 and 1979 than does its stationary counterpart, for both Miller and Anheuser-Busch; the predictions for Miller seem superior. In their entirety, the (nonstationary) transition probability estimates for 1977 are

$$\hat{P}(26) = \begin{array}{c} \\ \mathrm{B} \\ \mathrm{M} \\ \mathrm{O} \end{array} \begin{array}{c} \begin{array}{ccc} \mathrm{B} & \mathrm{M} & \mathrm{O} \end{array} \\ \begin{bmatrix} .9950 & .0046 & .0004 \\ .0000 & .9971 & .0029 \\ .0084 & .0522 & .9394 \end{bmatrix} \end{array},$$

whereas the stationary estimates are

$$\begin{array}{c} \\ \mathrm{B} \\ \mathrm{M} \\ \mathrm{O} \end{array} \begin{array}{c} \begin{array}{ccc} \mathrm{B} & \mathrm{M} & \mathrm{O} \end{array} \\ \begin{bmatrix} .9609 & .0391 & .0000 \\ .0000 & 1.0000 & .0000 \\ .0134 & .0000 & .9866 \end{bmatrix} \end{array};$$

[10] Again, analogy to linear regression theory suggests that under the null hypothesis,

$$\frac{(SSR_R - SSR_U)/[(R-1)^2]}{SSR_U/[(T-R)(R-1)]} \sim F_{(R-1)^2, (T-R)(R-1)}.$$

the corresponding market-share predictions for 1978 and 1979 are found in Table IV. (Compare Table IV with the independent published estimates in Table V.) Although both models predict a rise in Miller's market share, only the nonstationary model predicts the market-share increase for Anheuser-Busch.

TABLE IV
MARKET-SHARE PREDICTIONS FOR 1978 AND 1979 FROM THREE-STATE MARKOV PROCESSES

| *Year* | *Anheuser-Busch* | *Miller* | *Other* |
|---|---|---|---|
| Nonstationary Model | | | |
| 1978 | .2375 | .1869 | .5756 |
| 1979 | .2411 | .2175 | .5414 |
| Stationary Model | | | |
| 1978 | .2326 | .1634 | .6040 |
| 1979 | .2316 | .1725 | .5959 |

TABLE V
INDEPENDENT PUBLISHED ESTIMATES OF MARKET SHARES FOR 1978 AND 1979

| *Year* | *Anheuser-Busch* | *Miller* | *Other* |
|---|---|---|---|
| *Advertising Age* | | | |
| 1978 | .283 | .194 | .523 |
| 1979 | .304 | .245 | .451 |
| Kidder, Peabody & Co., Inc. | | | |
| 1978 | .2568 | .1932 | .5500 |
| 1979 | .2751 | .2149 | .5100 |

## C. *Some Observations*

From the two- and three-state results reported above, it appears that Miller's advertising effort considered alone may be more effective in changing its market share and affecting its transfer-purchase probabilities than the advertising efforts of Anheuser-Busch and other firms. These results suggest strongly that the $\alpha_{ij2}$'s are not the same for all $i$ and $j$, and are consistent with the argument that Phillip Morris's advertising campaign for its Miller acquisition has both a quantitative and a qualitative dimension (see, e.g., Mueller [16] for a discussion of Phillip Morris's cross-subsidization of Miller's advertising campaign).

## V. CONCLUSIONS

Employing a methodological approach which theoretically could be applied in the study of many socioeconomic processes, we have addressed the issue of advertising's effect on brand shift and market share. The results of this study indicate that brand-shift probabilities are indeed sensitive to intraindustry advertising differentials. Furthermore, Miller's advertising has exerted an especially significant effect on its transfer-purchase probability and market share.

This research does not mean to suggest that advertising or product differentiation *alone* is responsible for Miller's success, or the rising concentration in brewing (as Elzinga [4] carefully argues). It does, however, suggest a framework through which the dynamics of market-share change incorporate economic influences, the effect of which can now be estimated directly. It also clearly produces results consistent with the trend of increasing concentration in the brewing industry. Deterministic models may require strong behavioral assumptions, e.g., maximization of the net discounted value of profits for infinite time into the future (see Grabowski [5]); they may also require abstracting from differences in advertising effectiveness across firms. The stochastic process model offers a distinct alternative for estimating the effect of advertising on market share (and transition probabilities).

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## APPENDIX

The estimation procedure is to solve the following unconstrained nonlinear programming problem:

$$\text{(5)} \qquad \underset{\beta}{\text{Minimize}} \sum_{t=1}^{T} [x_*(t) - P'_*(\beta, t)x(t-1)]'[x_*(t) - P'_*(\beta, t)x(t-1)],$$

where (for *our* problem)

$\beta$ = a $(2R - 1)(R - 1) \times 1$ vector
$T$ = the number of transitions
$x(t)$ = an $R \times 1$ vector of entity proportions in each of the $R$ states at time $t$
$P_*(\beta, t)$ = an $R \times (R - 1)$ matrix with $(i, j)$th element $p_{ij}(\beta, t)$
$p_{ij}(\beta, t) = \exp[F_{ij}(\beta, t)]/\{1 + \sum_{m=1}^{R-1} \exp[F_{im}(\beta, t)]\}$
$F_{ij}(\beta, t)$ is given in (6) below
$x_*(t)$ = an $(R - 1) \times 1$ vector consisting of the first $R - 1$ rows of $x(t)$
$C'$ = the transpose of a matrix or vector $C$

(Deletion of the last row of $x(t)$ and defining $P_*(\beta, t)$ to have only $R-1$ columns avoid equation redundancy.)

It is convenient to segment $\beta$ as follows. For $i = 1, \ldots, R$ and $j = 1, \ldots, R-1$, let

$$\alpha_{ij1} = \beta_{R(j-1)+i}$$

and when $i \neq j$, let

$$\alpha_{ij2} = \beta_{R(R-1)+(R-1)(j-1)+k(i,j)}$$

where

$$k(i,j) = \begin{cases} i-1 & \text{if } i > j \\ i & \text{if } i < j \end{cases}.$$

Let $A_i(t)$ be the advertising expenditure of firm $i$ at time $t$. With this additional notation, define (for $i = 1, \ldots, R$, $j = 1, \ldots, R-1$, $i \neq j$, and $t = 1, \ldots, T$)

$$F_{jj}(\beta, t) = \alpha_{jj1},$$

(6) and

$$F_{ij}(\beta, t) = \alpha_{ij1} + \alpha_{ij2}[A_i(t-1)/A_j(t-1)].$$

To investigate the lagged effect of advertising on transfer-purchase probabilities, we compute

$$\text{(7)} \qquad \frac{\partial p_{ij}(t)}{\partial A_j(t-1)} =$$

$$\frac{\{-A_i(t-1)/[A_j(t-1)]^2\}\alpha_{ij2}\exp[F_{ij}(\beta, t)]\left\{1 + \sum_{m=1, m\neq j}^{R-1} \exp[F_{im}(\beta, t)]\right\}}{\left\{1 + \sum_{m=1}^{R-1} \exp[F_{im}(\beta, t)]\right\}^2}.$$

Since $A_i(t-1)$ is always (strictly) positive, $\partial p_{ij}(t)/\partial A_j(t-1)$ is positive if and only if $\alpha_{ij2} < 0$.

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